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Firm-level Evidence on the Labor Impacts of the East Asian Crisis

Mary Hallward-Driemeier
Bob Rijkers
Andrew Waxman

The World Bank
Development Research Group
Macroeconomics and Growth Team
&
Financial and Private Sector Development Network
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Abstract

In a crisis, do employers place the burden of adjustment disproportionately on female employees? Relying on household and labor force data, existing studies of the distributional impact of crises have not been able to address this question. This paper uses Indonesia's census of manufacturing firms to analyze employer responses and to identify mechanisms by which gender differences in impact may arise, notably differential treatment of men and women within firms as well as gender sorting *across* firms that varied in their exposure to the crisis. On average, women experienced higher job losses than their male colleagues *within* the same firm. However, the aggregate adverse effect of such differential treatment was more than offset by women being disproportionately employed in firms hit relatively less hard by the crisis. The null hypothesis that there were no gender differences in wage adjustment is not rejected. Analyzing how employer characteristics impact labor market adjustment patterns contributes to the understanding of who is vulnerable in volatile times.

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Firm-level Evidence on the Labor Impacts of the East Asian Crisis

Mary Hallward-Driemeier, World Bank[1]
Bob Rijkers, World Bank
Andrew Waxman, Cornell University

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1 Introduction

Using firm level data, this paper brings a new perspective to the question of who is vulnerable in volatile times. By focusing on gender differences in labor market impacts of crises, it illustrates how firm level data can benefit our understanding of who is at risk of losing their job and why. Existing studies that examine the distributional impact of crises almost exclusively rely on labor force and household level data, and consequently focus on labor supply. However, crisis induced changes in labor market outcomes are predominantly driven by changes in labor demand; labor supply responses are typically muted compared to the large contractions in labor demand that are a defining feature of crises (McKenzie, 2004; Thomas et al., 2002). In addition, micro-level studies of the distributional consequences of crises typically document significant heterogeneity in impact across broad sectors, yet fail to explore how such heterogeneity arises (see e.g. Fallon and Lucas, 2002). Moreover, they are not able to distinguish between different mechanisms that might explain why some workers are more vulnerable than others.

This paper helps redress these lacunae in the literature by examining how variations in firm responses during the East Asian crisis in Indonesia affected men's and women's relative employment vulnerability. It distinguishes between two previously unexplored mechanisms by which gender difference in impact might arise: differences in vulnerability could be the result of (i) *sorting* by gender into firms and occupations that differ in their vulnerability to crises, and/or of (ii) of *differential treatment* of men and women workers *within* firms.

That employers often treat men and women differently in the workplace is widely documented (see Altonji and Blank, 1999 for a review of the literature), and there is no shortage of rhetoric asserting that women are disproportionately at risk during volatile times (Grown, 2009; ILO, 2009; World Bank, 2009).[2] A common concern is that women get dismissed first and hired last (Oxfam, 2009, Seguino, 2009 and UNGEI, 2009). However, the limited empirical evidence on the labor market consequences of macroeconomic crisis is distinctly ambiguous as to whether women or men were more vulnerable (Houseman and Abraham, 1993, Lim 2000, Thomas et al., 2002), begging the question as to whether claims regarding women's vulnerability may be exaggerated.

Moreover, many of the assertions about women's relative vulnerability are predicated on the idea that employers discriminate against women. While we cannot prove or disprove the existence of

[2] See for example IFPRI, 2008; ILO, 2009; UNDP 2009; and World Bank, 2009; for discussions of the gender implications of the recent food, fuel and financial crises.

employer discrimination with our data, distinguishing between sorting and differential treatment helps assess the likely magnitude of employer discrimination. In the sorting channel, the role of gender is indirect. It could influence how men and women select into different firms, but is unlikely to explain the variation in the impact of the crisis across sectors, firms and occupations. Thus, if differences in impact are driven by heterogeneity in impact of the crisis across firms that differ in the gender composition of their workforce (sorting), then the likely role of discrimination is limited. By contrast, differential treatment is more direct; at issue is whether, confronted with a need to adjust, firms fired women first or cut their wages more. Concerns of discrimination are more pertinent if gender differences in impact are predominantly due to differential treatment. [3]

Distinguishing between sorting and differential treatment is particularly important since the effects of these mechanisms could offset each other, resulting in an underestimation of gender differences in impact in aggregate statistics. This could happen, for example, if women were working in large firms that were more resilient to the crisis, yet such firms were more likely to fire women. Indeed, this paper will demonstrate that when favorable sorting effects dominate, aggregate statistics can obscure unfavorable differential treatment.

Sorting and differential treatment are not only relevant for examining employment effects, they could also explain gender differences in wage adjustment. Sorting across firms and by occupation could affect whether women tend to face wage or labor adjustment. The extent to which an employer will respond by cutting wages rather than jobs is likely to be affected by the employer's characteristics, e.g. large firms and foreign owned firms tend to pay higher wages and may thus have greater flexibility to adjust wages than small firms already paying minimum wages (or even lower wages). There may also be differential treatment in wage adjustment within firms. If women are willing to accept greater wage cuts than men, there could also be a gender specific wage-employment tradeoff that may render their employment less vulnerable than men's. To our knowledge, this is the first paper to examine the wage-employment trade-off during a crisis and the only paper thus far to consider gender-differences in this trade-off.

The 1997-98 crisis in Indonesia provides a very relevant context to explore these issues. As in many other countries, gender norms in Indonesia prioritize men's employment over women's

[3] Differential treatment need not necessarily reflect discrimination per se since we cannot control fully for workers' individual characteristics. Conversely, in the presence of unobserved worker heterogeneity, it is theoretically possible, although unlikely, to have discrimination in the absence of observed differential treatment. In this paper, we control for time-invariant heterogeneity, which should capture the lion's share of unobserved differences between men and women. Moreover, we control for differences due to occupation and a wide range of firm characteristics.

(World Value Surveys, 1981-2009). Moreover, identification of firm responses is facilitated by the unexpected and "quasi-experimental" nature of the crisis and its severity, the high flexibility of Indonesian labor markets at the inception of the crisis, and the richness of the Indonesian manufacturing census dataset that not only provides detailed information on employer characteristics, but also gender disaggregated employment information by broad occupational categories (e.g white- and blue-collar workers).

The paper is organized as follows. The next section reviews the related literature and provides background information on the Indonesian crisis. The data are discussed in Section 3, which also provides a static overview of sorting patterns by gender and a bird's eye view of employment and wage adjustment during the crisis. Section 4 assesses sorting effects by examining whether firms that employed proportionately more women were more likely to exit and by analyzing how, among continuing firms, employment growth was correlated with the initial gender composition of the workforce. Next, Section 5 examines the evidence for differential treatment by assessing whether women were proportionately more likely to be fired than men in firms that adjusted the size of their workforce. It not only assesses whether men and women were treated differently at the firm-level, but also whether the crisis led to a change in such discriminatory practices/differential treatment of men and women within firms.[4] We therefore test how employment adjustment and wage setting during the crisis differed from non-crisis periods. Section 6 examines the evolution of the gender wage differences to assess whether women suffered more severe wage cuts than men, and the extent to which there is gender-specificity in employment-wage tradeoffs. Conclusions are presented in Section 7.

The appendix validates our firm-level analysis with an analysis of the evolution of wages and employment using data on manufacturing employees from SAKERNAS, a nationally representative labor force survey in Indonesia. We cannot tell from this labor force survey the characteristics of the employer, but the labor force data enable us to control for individual characteristics and the comparison with the firm-level data highlights the extent to which the patterns for the larger formal manufacturing sector were mirrored in the broader population of manufacturing workers.

[4] For example, if women are always more likely to be laid off when a firm is forced to downsize, one might also expect them to be more likely to be laid off in times of crisis. Alternatively, it could be the case that women are only discriminated against during times of extreme economic stress, but not in less turbulent times. An interesting question in this regard is whether women suffer disproportionately during the crisis relative to other times. We address this by examining how differential treatment during the crisis differed from differential treatment during non-crisis times.

Our main results can be summarized as follows: Gender differences arose both because of sorting and differential treatment and these worked in opposite directions for blue-collar workers but not for white-collar workers. Overall, sorting dominated – and was favorable to women's employment. Even though on average 46% of all manufacturing jobs were held by women over the period studied, women's net job losses accounted for only a third of all the jobs lost during the crisis. However, we find heterogeneity across occupations and types of firms. Blue-collar women suffered from unfavorable differential treatment in hiring and firing. By contrast, white collar women were more likely to retain their jobs than their male counterparts in such firms. While, the results show a large gender gap in wages, as well as substantial adjustment of wages during the crisis, there was little change in the wage gender gap during the crisis. These patterns are also mirrored in the labor force data analyzed in the appendix.

2 Previous Literature and Context

2.1 Related Literature

This study brings together and builds on two strands of literature that have evolved quite separately up until this point. While a large body of literature has examined gender differences in the labor market outcomes of crises and shocks, focusing especially on the supply side by concentrating on households and individuals,[5] few such studies have analyzed the role of employers. Conversely, the relatively small literature on firms' adjustment during crises has not explored whether firms' responses result in gender differences in wage and employment outcomes.

The literature on differences in men's and women's employment outcomes due to macroeconomic shocks has typically focused on changes brought on by either large-scale structural policy regime shifts (i.e., trade, privatization or financial liberalization) or by macroeconomic crises. [6]

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[5] A broader literature has explored the question of whether women bear the brunt of coping with economic shocks (see *inter alia*, Sen and Grown, 1987) and has yielded conflicting results; in some instances women and girls suffer more severe welfare effects than men and boys in terms of health and nutrition (see, for example, Behrman, 1998; and Dercon and Krishnan, 2002); in other instances, there are no gender differences in health and education outcomes (Levine and Ames, 2003) or in household income (Cunningham and Maloney, 2000); occasionally, men suffer more than women. For example, male headed households suffered larger income losses due to macroeconomic shocks in Peru 1980s (Glewwe and Hall; 1998).

[6] By examining how gender equity is affected by an economic crisis the paper also contributes to the growing literature on the relationship between gender equity and economic development (see e.g., Mammen and Paxson, 2000). While this literature has focused extensively on the link between growth and gender equity (see e.g., Klasen and Lamanna, 1998), it is not known whether gains in gender equity are wiped out during a crisis. For example, it is not known whether crises inflate or reduce preexistent gender pay differences.

Starting with the former, trade policy has been shown to have a substantial gender impact because of sorting by gender into exporting and non-exporting firms (Wood, 1991; Fontana and Wood, 2000; Ozler, 2000).[7] Rama (2002) considers the impact on women's employment and wages of a downsizing of state-owned enterprises in Vietnam using simulations based on the Vietnam Living Standards Survey and finds evidence for initial shedding of female workers (relative to men), but a reduction in gender pay gaps. Secondly, studies on gender differences resulting from macroeconomic crises have yielded conflicting conclusions about the responses of female workers relative to those of male workers (see e.g. Thomas et al., 2002).

Macroeconomic shocks have profound impacts on job flows and wages. Studies of manufacturing firms suggest that downturns due to business cycle fluctuations are typically associated with *excess* churning of firms and workers (Davis and Haltiwanger, 1992; Davis et al., 1996). Gallego and Tessada (2009) use sectoral panel data on job flows to examine the employment impact of sudden stops in Brazil, Chile, Colombia, and Mexico and find that crises are associated with lower job creation and increased job destruction and that these effects are heterogeneous across sectors and countries. Hallward-Driemeier and Rijkers (2010) show that in the case of Indonesia, the East Asian crisis was particularly destructive, of firms and jobs, with the usual correlation of productivity and survival breaking down during the crisis.

However, very few studies examine heterogeneity in impact by gender. We are aware of only two papers that look at gender differences in job creation and destruction focusing on firm-level employment dynamics. Houseman and Abraham (1993) examine differences in employment responses of men and women to changes in output in U.S. and Japanese manufacturing from 1970-1990. Although their study does not explicitly focus on employment adjustment in response to financial crises, their findings suggest that, relative to men, women bore a disproportionate amount of the burden of the labor adjustment to the oil crises of the 1970s. The second is Lim (2000), who examines employment patterns in the Philippines during the East Asian crisis and finds that male employment, having increased relatively faster in the lead up to the crisis then fell relatively more during the crisis. Yet these studies neither focus on both wage and employment adjustment nor distinguish between differential treatment and sorting.

A handful of studies have examined firm level employment and wage adjustment in the context of transition from state-controlled towards market-oriented economies (see Christev and

[7] It is worth noting, however that Kucera and Milberg (2000) examined the impact of trade expansion on women during a subsequent period, notably from 1978 until 1995, and did not find the same effect.

Fitzroy, 2002, Basu et al. 1997, Haltiwanger and Vodopivec, 2003) and in the context of trade liberalization (Revenga, 1997). However, the analyses are not disaggregated by gender and do not encompass a crisis period.

2.2 Context: The Indonesian Crisis

In the early 1990s Indonesia was at the height of an extended period of industrialization with manufacturing being an important engine of growth. (Dwor-Frecault et al. 1999). In the wake of the devaluation of the Thai Baht in July 1997, flows of foreign capital reversed drastically and the rupiah depreciated dramatically. Interest rates were raised to defend the currency, which exacerbated the decline in demand. Inflation rates, which had been 12%, shot up to close to 100% (Hill, 1999). GDP contracted severely in 1997, and fell by over 13% in 1998. Manufacturing output followed a similar pattern, though manufacturing growth had already suffered a mild dip in 1996. The drop in manufacturing output was large relative to the corresponding dip in GDP. Manufacturing, a sector with 46 percent women employees at the inception of the crisis, was one of the hardest hit sectors, making it an interesting case study to examine how employer responses to shocks affected men and women differently.

Estimates based on SAKERNAS indicate the spike in inflation facilitated real wage adjustment by reducing the relevance of nominal wage rigidities and minimum wages. Across the earnings distribution, both men and women experienced real wage cuts on the order of 40%. Yet the changes in total employment rates were relatively small: between 1997 and 1998, female employment in urban areas declined by 0.7% and male employment declined by 2.2% (Thomas et al. 2002). Labor force participation rates increased slightly more for women (1.5%) than men (1.4%) (see also Cameron, 2000).[8] Thus, the impact of the crisis on aggregate employment was muted (Fallon and Lucas, 2002).[9] Thomas et al. (2002) conclude that "the evidence suggests that short-run labor supply functions are fairly inelastic in Indonesia" (p. 177). If this is true, then a reduced form

[8] By contrast, their estimates based on the Indonesian Family Life Survey, a rotating panel of households, suggest that women's employment rates went up slightly, primarily because many women started working as unpaid family workers.

[9] The aggregate figures masks reallocation of employment across sectors, including a reversal of rural-urban migration patterns and an increase in informality. Both the industrial distribution of jobs shifted and patterns of job creation altered substantially: while manufacturing, construction and retail sectors had been growing rapidly in the lead-up to the crisis, that trend was reversed once it had begun.

approach to modeling relative employment changes will predominantly be identifying demand rather than supply effects.

3 Data and Descriptive Statistics

3.1 Data and Construction of Key Explanatory Variables

The Indonesian Manufacturing Census (*Statistik Industri,* or SI) contains longitudinal data on all Indonesian manufacturing establishments with more than 20 employees from 1993-2004. A number of features of this dataset render it very suitable for analyzing employer responses to crisis.[10] To start with, a major strength of the data is that it disaggregates firm-level employment into blue and white collar workers by gender. Thus, we can analyze the impact of the East Asian crisis on the employment prospects of men and women, allowing for further differentiation by the broad types of jobs held by men and women. Second and related, the data enable us to study *within-firm* gender differences in employment controlling for crude occupational category (i.e. differentiating between white and blue-collar workers). Third, by virtue of being a census the data allow us to assess to what extent job losses are driven by exit and reduced entry, and to what extent they are the result of employment adjustment by incumbents. Fourth, the data contain information on the wage-bill for blue- and white-collar workers separately. Following Amiti and Cameron (2007) and Harrison and Scorse (2009), we use this information to construct proxies for firm-level average wages for blue-and white collar workers by dividing those wage bills by number of blue- and white-collar workers respectively. Although these firm-level wage data are not gender-disaggregated, they help us assess gender differences in wage adjustment: following Hellerstein and Neumark (1995, 1999) we focus on the relationship between an establishment's average wage in an occupation and the share of female employees in that occupation.[11] Last but not least, the SI data contain information on a rich number of firm characteristics such as their sector, material input usage, output, capital stock,[12] ownership structure and whether or not they export (see appendix A for more details).

[10] The census treats establishments (plants) rather than firms as their units of interest. Less than 5% of establishments in the census are owned by multi-plant firms (see Blalock and Gertler 2004 for a thorough discussion).
[11] Ideally one would control for individual characteristics in testing for differential treatment effects. This information is not available, but we do include firm-level proxies for productivity and wages.
[12] Information on the capital stock was not recorded in 1996 and missing for some firms in other years. Concerns about possible selection biases have led us not to control for capital per worker in our regressions. The qualitative pattern of

Although it is rare to have data that both contain detailed information on firms and the workers they employ and are as comprehensive in coverage over the span of a crisis period, some limitations of the SI data should be kept in mind. First, the census only includes firms with 20 or more employees and is consequently not representative of the entire manufacturing sector. Moreover, 'entry' is defined as entry into the survey, which happens when a new establishment with 20 or more employees is started or when a pre-existing establishment passes this threshold. Similarly, 'exit' could be due to firm closure or dropping below the threshold. Any temporary lapses in reporting are not counted as exits.[13] Second, only net changes in employment at the firm-level are observed. This could understate the total amount of churning that takes place. Moreover, we do not know whether a reduction in firm size is accompanied by forced layoffs or achieved by natural attrition, that is, whether workers leave voluntary or involuntary. A related concern is that the definition of the employment variable changed slightly in 2001, with unpaid workers no longer being reported separately. However, the number of such workers had not been large, and the results are robust to excluding the period 2001-2004.[14] Moreover, our occupational categorization of workers in to blue-and white collar categories is crude and we may consequently fail to document subtler gender differences occurring at finer levels of aggregation.

In addition, the data do not contain information on individuals. For example, we have no information on the number of hours they work, their age, their tenure and experience. Consequently, changes on the intensive margin cannot be identified and it is not possible to control for workers' observed and unobserved personal characteristics.

Appendix B presents supporting evidence on the evolution of employment and wages in the broader manufacturing sector based on repeated representative cross-sectional labor force surveys, the SAKERNAS data. This analysis corroborates the results presented below and, suggests that in manufacturing, gender differences in crisis induced adjustment in hours worked were small.

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results, however, is robust to doing so, regardless of whether we impute the capital stock for all observations for which it is missing or whether we only use observations for which information on the capital stock is available. These results are omitted to conserve space, but available from the authors upon request.

[13] To examine the sensitivity of our results to the 20-person cut-off, we re-estimated the regressions presented in this paper using a 30, 50 and 100-people cutoff instead. The resulting pattern of results did not change qualitatively. Our results do not appear to be driven by the exclusion of firms with fewer than 20 employees. Results are not presented to conserve space, but are available from the authors upon request.

[14] These results are omitted to conserve space, but available from the authors upon request.

3.2 Sorting Patterns: Where Do Women Work?

Figures 1A-F plot kernel weighted polynomial bivariate regressions of the gender composition on firm-size, capital intensity, value-added per worker, wages and share of output exported, to document sorting patterns by gender. They reveal a very strong association between firm-size and the share of women in the workforce; small firms tend to be majority-male firms, while large firms tend to be majority female-firms. In part because they tend to be larger, firms that employ a larger proportion of women are also more likely to export. Yet, the share of women in the workforce is negatively correlated with capital intensity, value-added per worker and firm average wages, variables that are positively correlated with firm-size. Firms that use proportionately more blue-collar workers also employ proportionately more women (see figure 1F). The coexistence of these patterns points towards the importance of sorting by gender into different types of industries and firms.

Table 1 presents descriptive statistics by industry, ranking them on the basis of the share of women they employ. Industries which are highly labor-intensive pay relatively low wages and have a high export propensity, such as the tobacco and apparel industry, employ a lot of women. By contrast, capital intensive industries, such as transport and motor vehicles, basic machinery and equipment employ relatively few women. These patterns are by no means unique to Indonesia; very similar gender sorting patterns have been documented in a host of other countries (see e.g. World Bank forthcoming, Chapter 5 and the references therein). Possible explanations for these sorting patterns are that women have a comparative advantage in industries that require fine motor skills, women may be less likely to be involved in labor disputes which can be a particular concern in large firms, and that women may lack more favorable outside opportunities (relative to men).[15]

Given the sorting of women into larger firms, we will present weighted regressions. Unweighted statistics reflect average outcomes for firms, while weighted statistics reflect outcomes for workers, better capturing average outcomes for women.[16] While women represent 46 percent of blue-collar workers in manufacturing, only about a third of all white-collar workers are women.

[15] For example, according to study of manufacturing firms in Bangladesh (Paul Majumder and Begum, 2000, p3) " *women sort into garment firms because a) women are more patient and nimble; (b) women are more controllable than men; (c) women are less mobile and less likely to join a trade union; and (d) women can do better in sewing because this job coincides with their traditional jobs. Most of the garment employers interviewed, reported that benefits arising from these qualities amply compensate the cost of employment of women in terms of maternity leave, high absenteeism, and other factors".*

[16] The results obtained using unweighted regressions are in general very similar to those obtained using weighted regressions. An interesting difference in weighted and unweighted pertains to differential treatment, the evidence for which is somewhat stronger in unweighted regressions. This suggests that differential treatment is more prevalent in small firms. Results are not presented to conserve space, but available from the authors upon request.

However, white-collar workers account for no more than 15% of all manufacturing workers. Whereas the female share of blue-collar workers varies from 11 percent in motor vehicles to 72 percent in tobacco, there is much less variation in the share of women in white collar jobs across industries.

White-collar workers are typically better educated than blue-collar workers (see also Lipsey and Sjoholm, 2004 and Harrison and Scorse, 2009). Gender differences in education vary within occupation: Male blue-collar workers are on average slightly better educated than female ones, whereas female white-collar workers tend to be somewhat better educated than male white-collar workers.

3.3 A Bird's Eye View of Aggregate Adjustment Patterns

3.3.1 Employment Adjustment

Following Dunne et al. (1989), figures 2 present descriptive statistics on job flows separately by gender, with figure 2A documenting patterns for blue-collar workers and figure 2B documenting job flows of white collar workers. For both groups of workers, all measures of job flows become smaller in magnitude over time; both excess reallocation and net job creation diminished over time, consistent with declining labor market flexibility over time due to more stringent labor regulation.[17] The crisis is associated with a large drop in manufacturing employment opportunities with white-collar workers suffering more severe job losses than blue-collar workers. Women seem to have fared slightly better in both occupational categories. While the difference in employment adjustment between men and women is small relative to total men and women employed, it is a substantial proportion of the total adjustment undertaken. With women representing 46% of the workers, they only suffered 34% of the net employment losses.[18] Thus the small gender differences in employment impact are in part attributable to men being more likely to be having white-collar jobs (sorting by occupation) and such jobs being more likely to be destroyed, and in part due to proportionately smaller reductions in the number of female jobs lost relative to men in both occupations.[19]

[17] After the fall of Suharto, policy priorities shifted, with greater labor protections put into place (see World Bank, 2010), Hallward-Driemeier et al. 2010b.

[18] These differences are also substantial in terms of the number of people affected; A one percentage adjustment in the number of jobs for men (women) amounts to roughly fifteen thousand jobs.

[19] Also note that the crisis led to a spike in excess churning driven by a simultaneous increase in job destruction and a decrease in job creation. Nonetheless, the fact that, even during the crisis, overall job creation rates were as high as 13.8% in 1997 and 16.3% in 1998 is indicative of high heterogeneity in firm performance.

Figures 3A and 3B further distinguish between job creation and destruction due to adjustment by incumbents, exit and entry of new firms, separately for blue- and white-collar workers respectively. Gender differences are largely driven by job destruction by incumbents, which were most severe for men and account for the bulk of the decline in net job creation. Overall then, these findings suggest that while employment adjustment patterns for men and women look broadly similar, women fared better.[20]

3.3.2 Wage Adjustment

Figures 4A and 4B plot the evolution of firm-level average wage growth for blue-collar and white-collar workers. To assess how wage adjustment varied with the gender composition of the workforce, separate plots are drawn for firms where women comprise less than one third (male dominated), one to two thirds (gender neutral) and more than two thirds (female dominated) of the workers. For comparison, plot 5C depicts the evolution of inflation. Clearly, the spike in inflation that was precipitated by the devaluation of the Rupiah eroded real firm-average wages which declined by 22% for blue-collar workers in 1998 and by 28% for white-collar workers on average (see figures 4A and 4B). White-collar workers thus suffered both larger employment losses and larger wage cuts. After the crisis, real wage growth resumed. Consistent with the macroeconomic trends described in Section 2, firms coped with the crisis both by cutting employment and reducing real wages.

Plots 4A and 4B show that average firm wage adjustment was not very strongly correlated with the gender composition either for blue-collar or white-collar workers. However, the averages presented here mask significant heterogeneity within and between firms. The next three sections analyze this heterogeneity in more depth.

4 Sorting

At this juncture it is useful to provide a roadmap of the micro-analysis that lies ahead. Section four assess the importance of sorting in employment adjustment by analyzing the relation

[20] In interpreting these findings it is important to bear in mind that we do not have any information whether or not separations were voluntary. For example, we cannot rule out the possibility that men who left manufacturing did so because they were able to take up more lucrative income earnings opportunities. However, this scenario seems somewhat in plausible in view of the fact that earnings declines were widespread and not confined to the initially affected sectors (Fallon and Lucas, 2002).

between the gender composition of the workforce and firm-survival (section 4.1) and employment growth of continuing firms (4.2). Gender patterns could well differ between survival and growth, particularly if women are predominantly in larger firms that are more likely to survive, but then grow slower on average. Section 5 then turns to the question of differential treatment. Changes in wages are then examined in Section 6.

4.1 Firm Survival

Figure 5A tracks the evolution of entry and exit rates while Figure 5B plots how the gender composition of the workforce of exiting firms evolved over time. The crisis is clearly associated with a spike in exit rates.[21] Moreover, the average share of female workers in exiting firms dropped slightly during the crisis, consistent with the figures presented in Section 3.

4.1.1 Econometric Framework

To model firm survival, a discrete-time proportional hazards survival model is used, which is estimated by means of logistic regression.[22] Following Cox (1972) we assume that the hazard rate, $\lambda(t_i|\boldsymbol{x})$ of a firm characterized by covariates x_i, is proportional to a baseline hazard λ_0:

$$\frac{\lambda_t(t_i|x_i)}{\lambda_0(t_i|x_i)} = \exp\left(\beta_{x_i}{}'\boldsymbol{x_i}\right) \qquad <=> \quad log\ \lambda_t(t_i|\boldsymbol{x_i}) = log\ \lambda_0(t_i|\boldsymbol{x_i}) + \beta_{x_i}{}'\boldsymbol{x_i}$$

This choice of the hazard function implies that each regression coefficient β_{x_i} measures the proportionate change in hazards associated with absolute changes in the corresponding covariate $x_i \in \boldsymbol{x_i}$. For example, if the share of women is not correlated with an increased risk of exit then $\beta_f = 0$, and $\beta_f > 0$ ($\beta_f < 0$) is associated with an increased likelihood of exit (survival).

[21] There is another spike in exit in 2000, which is most probably driven, at least in part, by changes in survey design (see the Appendix for a discussion of this issue).

[22] In discrete time, the survival function $S(t_i|\boldsymbol{x})$, which measures the probability that for a firm *i* with a vector of characteristics, $\boldsymbol{X}$, survival time, T, is at least t_i, can be expressed as the product of the complements of period specific hazard rates: $S(t_i|x) = \Pr\left(T \geq t_i|\boldsymbol{x}\right) = \prod_{i=1}^{t}(1-\lambda(t_i|\boldsymbol{x}))$. The $\lambda(t_i|\boldsymbol{x})$ refer to the period specific hazard rates, which measure the conditional probability of exiting at time t_i conditional on surviving until time t_{i-1} and are defined as $\lambda(t_i|\boldsymbol{x}) = \Pr\left(t_{i-1} \leq T < t_i|T \geq t_{i-1}, x_{i-1}\right) = 1 - \frac{f(t|\boldsymbol{x})}{F(t|\boldsymbol{x})} = \frac{f(t|\boldsymbol{x})}{S(t|\boldsymbol{x})}$, where $F(t|\boldsymbol{x})$ is the cumulative density function of time at risk given the vector of characteristics $\boldsymbol{x_i}$.

To examine how the determinants of firm-survival varied during the crisis, these explanatory variables are interacted with a dummy for the crisis period. Adding an error term $\boldsymbol{v_{it}}$ yields the estimable equation:

(1) $$log\ \lambda_t(t_i) = log\ \lambda_0(t_i) + \beta_{x_i}{}' \boldsymbol{x_i} + \beta_{Crisis*x_i}{}' Crisis * \boldsymbol{x_i} + \beta_{Crisis_i}{}' Crisis + \boldsymbol{v_{it}}$$

where *Crisis* is a dummy variable for 1997 and 1998. If $\beta_f + \beta_{Crisis*f} > 0$ (< 0), then the likelihood of exit increases (decreases) in absolute terms with the share of women employed during the crisis – ceteris paribus. The crisis interaction thus helps us assess whether firms employing more women were more likely to exit during the crisis relative to other periods. This testing strategy is very general as we allow all parameters of the hazard function to vary, thereby minimizing parameter constancy restrictions.

4.1.2 Results

Table 3 present the results of our survival analysis and presents four specifications. The basic specification presented in column 1 only includes controls for the gender composition of the workforce whereas the second specification, which is presented in column 2, includes additional controls for a rich set of firm characteristics including firm size, the skills composition of the workforce (proxied by the "unskilled ratio"), whether or not the firm exports, firm age, foreign as well as government ownership and productivity, proxied by value added per worker, and sector-year dummies, which not only control for sectoral differences but also for sector-specific time-varying shocks. The first specification captures unconditional gender differences, while the second column documents to what extent gender differences are due to gender per se and to what extent they are the result of differences in firm characteristics correlated with the gender composition of the workforce. A comparison of the results from the two columns is thus revealing about the importance of sorting. Columns 3 and 4 replicate these specifications, but allowing for gender differences by occupation instead of the total gender composition of the workforce.

The first two specifications demonstrate the importance of sorting. The baseline specification reveals that during non-crisis times the relationship between firm exit and the gender composition of the workforce is significant and negative, and that during the crisis this association becomes significantly more negative; during the crisis women were thus much less likely to lose their jobs due to firm exit. The second specification shows that this is entirely a sorting effect; once other firm characteristics are accounted for, both the coefficient on the gender composition of the workforce and its interaction with the crisis become insignificant. The regression furthermore confirms that white collar workers were hit harder; the coefficient on the interaction term between the crisis and the unskilled is significantly negative, suggesting that employing proportionally more white-collar workers is associated with a heightened exit propensity.

The third and fourth columns decompose the gender effects by broad occupational category and show that the relatively favorable sorting effect for women during the crisis is confined to blue-collar workers only; they are significantly less likely to lose their jobs during crises ceteris paribus, whereas white collar women were not less exposed than white-collar men.[23] Note, however, that during non-crisis times, women in white collar jobs are at a marginally lower risk of losing their jobs once we condition on the gender and occupational composition of the workforce, as is evidenced by the negative coefficient on the proportion of white collar workers that are women.[24]

Overall, these findings highlight the importance of sorting; gender differences in employment losses due to firm exit are driven by sorting into firms and occupations that differed in their vulnerability to the crisis.

4.2 Firm Growth

Figures 6A and 6B plot bivariate, locally-weighted polynomial regressions of the change in the log of the number of blue- and white-collar workers, respectively, on the gender composition of the workforce in the prior period, both during the crisis and during non-crisis years. Firms on average shed labor during the crisis, as the crisis lines lie substantially below the non-crisis lines in both plots. That white-collar workers were hit much harder by the crisis than blue-collar workers is evident when one notes the different scaling of the axes of the two figures.

Employment growth of blue-collar workers (Figure 6A) is very mildly negatively correlated with the share of women in the workforce during non-crisis years and not correlated with the gender composition during crisis years. Figure 6B provides evidence that the relationship between the proportion of female white-collar workers and the change in the total number of white-collar

[23] The interaction term for the female share of white collar jobs and the crisis is only marginally insignificant – and positive. Together it would offset the average benefit that a higher share of female white collar employees had on the probability of exit. As such, the evidence points to female white collar employees sorting into firms that were somewhat more vulnerable during the crisis.

[24] Note, however, that the "unskilled ratio", which is positively correlated with the share of women in firms white and blue collar' workforces, is also significantly positively correlated with firm exit.

workers is not linear. Nevertheless, firms with a higher share of female white-collar workers generally grow slightly faster.

4.2.1 Econometric Strategy

To model labor demand we follow Svejnar and Basu (1997) and start from the general model: $L_{it} = F(Y, W, O)$ where labor usage of firm i is a function of firm output, *Y*, wages, *W*, and other firm characteristics O. This conceptualization of labor demand corresponds to a cost minimization framework in which enterprises set employment in response to exogenously given demand and input and output prices. To allow for the fact that adjustment is not instantaneous due to the presence of adjustment costs we introduce lags of both labor usage as well as the principal explanatory variables. Writing the model in log-linear form and adding an error term yields the estimable equation.

$$lnL_{it} = \alpha_{Y_t} lnY_{it} + \alpha_{Y_{t-1}} lnY_{it-1} + \alpha_{W_t} lnW_{it} + \alpha_{W_{t-1}} lnW_{it-1} + \alpha_{O_t} O_{it} + \alpha_{O_t-1} O_{it-1} \alpha_{L_{t-1}} lnL_{it-1} + \varepsilon_{it}$$

As pointed out by Basu et al. (2005) this specification nests a number of different specifications, including i) a completely static model ($\alpha_{W_{t-1}} = \alpha_{Y_{t-1}} = \alpha_{X_t-1} = \alpha_{L_{t-1}} = 0$), ii) a partial adjustment model ($\alpha_{W_{t-1}} = \alpha_{Y_{t-1}} = \alpha_{X_t-1} = 0$), and iii) a first-difference fixed effect mode ($\alpha_{W_t} = \alpha_{W_{t-1}}, \alpha_{Y_t} = \alpha_{Y_{t-1}}$, $\alpha_{Ot} = \alpha_{O_t-1}$, and $\alpha_{L_{t-1}} = 1$).

While our aim is to provide descriptive evidence, the above labor demand model can be derived from an underlying optimization framework; as shown by Nickell (1987), the specification is the solution to a cost-minimization problem of a firm with a Cobb-Douglas production function that faces exogenous output constraints and quadratic adjustment costs in labor and capital. Thus, the specification has a structural interpretation.

Since we are interested in examining the impact of the crisis and whether or not firms that employed more women experienced proportionately more employment losses, we interact the explanatory variables with a crisis dummy. In addition, following Christev and Fritzroy (2002) we impose the constraint that $\alpha_{Ot} = 1$ to avoid multicollinearity and separate firm characteristics into the gender composition of the workforce, *f*, and other firm characteristics *X*. To examine whether or not workers in firms that employ more women are more likely to be laid off, we thus estimate the change in labor demand, separately for blue-collar and for white-collar workers, using the following autoregressive distributed lag specification:

(2)
$$\Delta lnL_{it} = \beta_{\Delta Y_t} ln\Delta Y_{it} + \beta_{Crisis\Delta Y_t} Crisis * ln\Delta Y_{it} + \beta_{Y_{t-1}} lnY_{it-1} + \beta_{CrisisY_{t-1}} Crisis * lnY_{it-1} + \beta_{\Delta W_t} ln\Delta W_{it} + \beta_{Crisis*\Delta W_t} Crisis * ln\Delta W_{it} + \beta_{W_{t-1}} lnW_{it-1} + \beta_{Crisis*W_{t-1}} Crisis * lnW_{it-1} + \beta_{fit-1} * f_{it-1} + \beta_{Crisisfit-1} * \mathrm{Crisis} f_{it-1} + \beta_{Xt} * X_{t-1} + \beta_{Crisis*Xit-1} Crisis * X_{it-1} + \beta_{\Delta Xit} * \Delta X_{it} + \beta_{Crisis*\Delta Xit} * (Crisis * \Delta X_{it}) + \beta_{Lit-1} * L_{it-1} + \beta_{Crisis*Lit-1} * (Crisis * L_{it-1}) + \beta_{Crisis} Crisis + \varepsilon_{it}$$

If firms that employ proportionately more women grow faster (slower) than those that do not, then $\beta_{fit-1} > 0$ (< 0). The coefficient on the interaction $\beta_{Crisis.fit-1}$ tells us whether employment losses for these firms were larger in the crisis relative to other periods. The "absolute" adjustment during the crisis is given by $\beta_{fit-1} + \beta_{Crisisfit-1}$.

A number of econometric issues have to be addressed. First, to minimize the role of outliers we use the "Haltiwanger growth measure", i.e. the change in employment over the average firm-size for two periods: $H = (L_t - L_{t-1})/[0.5*(L_t + L_{t-1})]$, thus bounding the growth rate between -2 and +2. Second, serial correlation in the error term in conjunction with the presence of lagged size as a regressor will cause OLS estimates of equation 2 to be biased. Fixed-effects estimates, too, will be biased because the transformed residual is by construction correlated with the transformed lagged dependent variable, the Nickell bias (Nickell, 1981), but in the opposite direction. Comparison of OLS and FE estimates will thus provide us with a confidence interval within which we expect the true parameter estimates to lie (Bond, 2002).[25]

4.2.2 Results

Table 4 presents regressions of growth in the employment of blue- and white-collar workers within surviving and new firms. We first focus on blue-collar workers as they account for the bulk of manufacturing employment. The basic specifications are presented in columns 1 and 2, with additional firm controls added in columns 3 and 4. Columns 1 and 3 provide the OLS results, with columns 2 and 4 the FE specifications.

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[25] In principle this endogeneity issue could be tackled by using GMM methods, however, the difference GMM estimator suffers from weak instrument bias if the underlying series are highly persistent—as is the case with the employment series under consideration. While the Systems GMM estimator could, in principle, remedy this problem, the additional moment conditions it invokes rely crucially on a mean stationarity assumption, which is a particularly unappealing assumption in the context of a crisis (see Roodman, 2006 for a discussion). We therefore eschew this approach.

The association between the gender composition of the workforce and labor growth is not significant in the OLS specification, but is significant in the FE and negative. It implies moving from having no women to exclusively employing women would reduce the estimated average growth rate by up to 4.2% (column 2). The crisis did lead to a statistically significant change in this association, with the interaction term having a significant positive coefficient; during the crisis firms with more women employees grew relatively faster (or shrank less rapidly). The effect is similar for both the OLS and FE regressions.

Columns 3 and 4 again demonstrate the importance of taking sorting across firms into account. In these regressions, the full set of firm characteristics is included, as are their interactions with the crisis dummy. The coefficient on the share of female workers is significantly associated with employment growth, but the sign has now flipped and is *positive*. Controlling for these firm characteristics, a greater share of female employees is associated with higher employment growth – with the effect significantly larger and still positive during the crisis. The overall magnitudes of the effects are similar between columns 3 and 4, with the crisis effect somewhat larger in the FE specification. In understanding the change in sign of the coefficient on the gender share, controlling for firm size is of particular importance. Underlying the negative unconditional correlation between the proportion of women and firm growth is that women are more likely to work in larger firms, which tend to grow less quickly than small firms.

The results on other firm characteristics generally accord with intuition: firms which paid higher wages last period and firms that increased wages, grow less quickly. This is consistent with the existence of a trade-off between wage and employment growth. Interestingly, during the crisis these associations are even stronger, hinting at the possibility of a stronger trade-off between wage and employment adjustment. Of course, one has to bear in mind that wages are potentially endogenous.[26] In addition, exporters were less likely to shed blue-collar workers during the crisis, presumably because they benefited from favorable exchange rate movements.

In Table 4, columns 5-8 present the results for white-collar workers. The most striking finding is that employment growth of white-collar workers is always positively correlated with the share of female white-collar workers, but the crisis interaction term is never significant at the 5% level, though it is negative and significant at the 10% level in both the OLS and FE specifications

[26] Since the firm level average wage is a composite of men's and women's wages, these specifications also go a long way towards controlling for potential differences in employment growth due to possible gender differences in willingness to absorb real wage reductions, as these should be reflected in firm-level average wage changes.

that include a full set of controls. The size of the coefficients on the gender share do fall in columns 7 and 8; sorting helps explain some of the favorable association with female white collar workers.

In summary, the gender composition of the workforce is correlated with growth in employment. For blue-collar workers the proportion of women in the workforce becomes less negative – and even turns positive – as we include firm controls, attesting to the importance of sorting. Our results are furthermore consistent with the existence of a(n exacerbated) tradeoff between wage and employment adjustment.

5 Differential Treatment

Figures 7A and 7B plot the change in the share of women against total labor growth of the relevant category of workers for both blue- and white-collar workers, separately for both the crisis and non-crisis periods. The mild positive association between an increase in the proportion of female blue-collar workers in the firm and growth in the number of blue-collar workers suggests that as firms expand their blue-collar workforce they hire proportionately more women (consistent with figure 1A). However, proportionately more women are laid off when they contract. In other words, women's employment appears more pro-cyclical than men's employment. The figures do not suggest that this relationship was distinctly different during the crisis.

For white-collar workers, these relationships are slightly different. Figure 7B portrays a very small negative correlation between changes in the number of white-collar workers and the proportion of white-collar workers who are women during non-crisis times. Again, the crisis association does not seem markedly different from the association observed in less volatile times. While the plots are striking as the observed unconditional correlations are small, the magnitude of the overall effects of differential treatment could be larger given the increased changes in employment during the crisis.

5.1 Econometric Approach

To assess whether men and women with the same occupation are treated differently within the same firm, the change in the proportion of women in firm i's workforce is regressed on the total amount of labor adjustment that firm i undertakes and a set of explanatory variables. The most general specification distinguishes between positive and negative changes in total employment and interacts all explanatory variables with a crisis dummy:

$$\Delta f_{it} = \delta_{\Delta lnLt}{}^{+}_{-} \Delta lnL_{it}{}^{+}_{-} + \delta_{Crisis\delta_{\Delta lnLt}{}^{+}_{-}} Crisis * \Delta lnL_{it}{}^{+}_{-} + \delta_{Xit} X_{it-1} + \beta_{CrisisXit-1} Crisis * X_{it-1} + \beta_{Crisis} Crisis + m_{it}$$

Where the + and – indicate positive and negative employment changes respectively, thus enabling us to test for potential asymmetries between net hiring and firing, and m_{it} is a random error term. The explanatory variables X_{it-1} allow us to control for, and characterize, firm heterogeneity.

5.2 Results

The results for blue-collar workers are presented in Table 5, columns 1-4. All specifications examine the impact of changes in the number of blue-collar workers (measured in logs), entered separately as either expansion (or no growth) and contraction, and their interactions with a crisis dummy. Overall, the regressions demonstrate that growth in the number of blue-collar workers is correlated with a small increase in the share of women, while a contraction is associated with a symmetric reduction in the proportion of women. This finding is consistent with the positive association between firm-size and the share of blue-collar workers that are women documented in Section 3. The crisis did not induce a change in these relationships relative to other periods, except in the case of the FE specification with firm controls, where the effect was doubled during the crisis. As average employment growth was negative during the crisis, female blue-collar workers thus suffered comparatively larger employment losses as a result of differential treatment than male blue-collar workers did. However, it should be kept in mind that the magnitude of these effects are very small; increasing (reducing) the number of blue-collar workers by 10% is associated with a 0.07% increase (decrease) in the share of blue-collar workers that are women.[27]

The results for white-collar workers are also reported in Table 5, in columns 5-8. They reveal a negative association between growth in the number of white-collar workers and the share of female white-collar workers. The finding is robust to controlling for the full set of controls (columns 7 and 8). This is consistent with Figure 1A; except for the very largest firms, the association between size and the share of female white-collar workers trends downwards. When

[27] If these specifications are run unweighted, the associations between changes in size and changes in the gender composition gain significance, suggesting that changes in the gender composition of the workforce associated with changes in firm-size are largest in small firms.

firms expand the number of white-collar workers, they are likely to hire proportionately more men (relative to the prevailing gender composition of the white-collar workforce). Yet, when they reduce the number of white-collar workers, women are more likely to preserve their jobs (in relative terms). During the crisis, the impact of differential treatment in expanding firms is significantly muted, such that women would be less disadvantaged. For contracting firms, the crisis serves to reinforce the pattern that women would be more advantaged, although the magnitude of this crisis effect is small (and not always significant). Thus the effect of the crisis on female white collar workers was less harmful on average than the effect on their male colleagues.

In summary, there is some evidence for differential treatment. Intriguingly, the impact of differential treatment varied by occupation. Firms that contracted their blue-collar workforce fired proportionately more women, while firms that reduced their white-collar workforce fired proportionately more men. These effects were then reinforced during the crisis. Only the effect of expanding firms hiring fewer white collar men was muted during the crisis. One possible explanation for these findings is that firms were more likely to retain their most skilled employees, which, in the case of blue-collar workers were men, and in the case of white-collar workers were women. However, the magnitude of these differential treatment effects is small.

6 What Happened to Relative Wages? The Evolution of Gender Pay Differences

6.1 Econometric Approach

The comparatively favorable employment consequences of the crisis for women might be due to a change in the relative price of female labor versus male labor. Following Amiti and Cameron (2007) and Harrison and Scorse (2010) we regress the logarithm of these firm-level average wages by occupational category, $\ln(\overline{w}_{it})$, on the share of women, $fempsh_{it}$, and a vector of firm characteristics, X_{it}. We also interact these explanatory variables with a crisis dummy to assess whether and if so, how, wage setting during the crisis differed from normal times. That is, we estimate regressions of the form:[28]

[28] Dynamic wage equations that model wage-growth as a function of the gender composition of the workforce and other firm characteristics yield very similar results and are available from the authors upon request.

$$\ln(\bar{w}_{it}) = \gamma fempsh_{i,t} + \gamma_{\text{Crisis}}\text{Crisis} * fempsh_{i,t} + \beta_{\text{X}}X_{it} + \beta_{\text{Crisis}}\text{Crisis} * X_{it} + \beta_{Crisis}Crisis + u_i + v_{it}$$

where u_i is a time invariant and unobserved firm-specific fixed effect and v_{it} is an error term that is assumed to have zero mean and finite variance. A correlation between u_i and any of the regressors can cause OLS estimates to be biased. Our preferred estimator is therefore the Fixed Effects estimator; wage adjustment is identified on the basis of within-firm changes, and any potential bias due to unobserved time-invariant firm characteristics is removed. The coefficient on the share of women, γ, measures the differences in pay between men and women, conditional on observable firm characteristics as captured in X_{it}, and can be considered a crude proxy for the gender pay gap (see Hellerstein and Neumark, 1999) [29] under the assumption that workers of the same gender and with the same occupation earn the same wages.[30] To see this note that increasing the proportion of women from 0 to 1 would increase estimated log firm average wages by γ, or approximately γ %.

6.2 Results

Results are presented in Table 6. Again we compare the extent to which unconditional gender pay differences are associated with gender, *per se*, and to what extent they are accounted for by other factors. Columns 1 and 2 present our baseline specifications which only control for the gender composition of the workforce and thus provide insight into average unconditional gender differences, while columns 3 and 4 add a full set of explanatory variables, including firm-size, ownership dummies, value-added per worker, firm age, whether or not the firm exports, sector-year dummies, as well as measure of real minimum wages. The share of women is strongly negatively correlated with firm-level average wages, which is indicative of gender-pay differences (column 1). Controlling for firm-fixed effects reduces gender pay differences in half (column 2). Adding additional firm controls[31][32] further reduces the pay gap. Both these results demonstrate that men sort

[29] It has to be interpreted with caution since we cannot identify changes in the earnings distribution within firms and lack controls for time-varying and time-invariant characteristics of the workforce such as education and hours worked.
[30] The empirical specifications below relax this assumption by including province, region and sector-yeardummies, as well as controls for a rich set of firm characteriscs.
[31] These results are robust to controlling for capital per worker and controlling for productivity using measures of TFP rather than value-added per worker. Using these controls leads to a substantial reduction in sample size since information on capital is missing for a large number of firms and missing altogether in 1996 – which is why we prefer

into higher paying firms (column 3). Nonetheless, the fact that the gender premium is robust to controlling for time-invariant characteristics suggest that gender pay differences are not exclusively driven by sorting (column 4). Also note that the Fixed Effects and OLS estimates become very similar.

The crisis does not have a significant impact on gender pay differentials, regardless of whether we control for firm characteristics and/or firm-fixed effects. Thus, on average, women blue-collar workers did not experience larger or smaller wage losses than men in the crisis.

Results for white-collar workers are presented in columns 5-8 and again provide evidence for gender pay differentials favoring men, which are in large part due to sorting.[33] Crucially, the crisis does not significantly impact gender pay differences for white-collar workers either. In summary, wage losses were severe, but gender differences in wage adjustment were statistically negligible for both blue- and white-collar workers.

7 Conclusion

This paper has examined whether and to what extent firm responses to the East Asian crisis affected men and women's manufacturing employment and wages differently. Adjustment took place via a reduction in employment as well as a drastic decline in real wages, facilitated by high inflation. Gender differences in the impact of the crisis on manufacturing employment were small in absolute terms but substantial relative to the total amount of adjustment undertaken. Women's employment was less affected than men's employment, but women did not suffer larger wage cuts, even though our regressions suggest that the wage-employment tradeoff became more acute during the crisis.

The comparatively favorable employment consequences for women were predominantly due to sorting by gender into firms that were less vulnerable to the crisis. Most saliently, women's employment was concentrated in larger firms and firms that export. Such firms exhibited lower net

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specifications that exclude capital and use value-added per worker as our proxy for productivity. Nonetheless, results are available from the authors upon request, but omitted to conserve space.

32 Size is positively correlated with average firm-level wages in the OLS specifications while the impact of size in the fixed effects regressions is negative, perhaps because of measurement error (recall that the average wage variables is computed by dividing the total wage bill by the total number of employees).

33 The OLS estimate of the unconditional gender wage gap is large and statistically significant; women white-collar workers on average earn 22% less than male white-collar workers. The FE regression shows the opposite pattern; that within firms a higher share of women white collar workers is associated with higher wages. Adding firm controls affects the results; The OLS coefficient remains significant but is again much smaller. In the FE regression there is now no effect; the unconditional effect can be explained by sorting rather than by the gender composition per se.

employment losses. Although they were less likely to grow conditional on survival, larger firms were more likely to survive, such that, in net terms, larger firms shed less labor. Part of the sorting by gender also occurred along occupational lines.[34][35] That is, men also suffered more because they were more likely to be working in white-collar jobs, which were more vulnerable.

Differential treatment effects varied by occupation as well as firm size, and are obscured in aggregate statistics, which reflect the combined effect of sorting and differential treatment. When firms contracted, the share of blue-collar workers that were female decreased somewhat, while the share of white-collar workers that were female increased. For blue-collar workers, these relationships were slightly exacerbated during the crisis, such that women were more likely to suffer job losses than their male colleagues within the same firms. By contrast, differential treatment protected white collar women relative to their male colleagues. The crisis mildly (but not consistently statistically significantly) reinforced the pattern that women in white collar jobs are more likely to be spared when firms contract. Moreover, they benefitted significantly more from firm expansion than during non-crisis times. However, these differential treatment effects were very small. That is, even with evidence of differential treatment, the claims about crises inducing a vast surge in employer discrimination discussed in the introduction appear grossly exaggerated; if discrimination was really driving women's vulnerability, one would have expected to see much more dramatic differences.

This paper has documented heterogeneity in the relative importance of these channels across different types of firms and workers in different occupations. It has, moreover, demonstrated how aggregate statistics may mask countervailing transmission channels. Examining to what extent this is the case using matched employer-employee data, and assessing to what extent our findings generalize to other sectors and other crises, are promising areas for further research. Overall, this paper underscores the importance of analyzing and documenting heterogeneity in firms' adjustment to major macroeconomic shocks in understanding the distributional impacts of crises.

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[34] However, white-collar workers comprise only a small share of all manufacturing employment, such that aggregate gender differences in crisis impact associated with sorting into different occupations are small.

[35] We also examined whether or not the fact that white-collar workers were hit harder by the crisis was due to within-firm adjustment in the occupational structure of the workforce (for example, one might speculate that firms have a proclivity to fire disproportionately more white-collar workers when they are confronted with a shock) but did not find any evidence that this was the case. Results are omitted to conserve space but available from the authors upon request.

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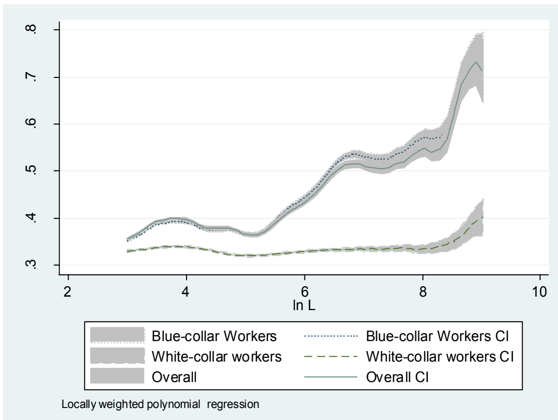
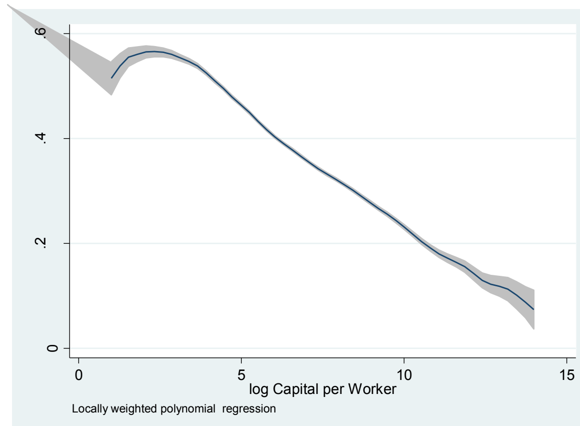
Figures 1A-K: Bivariate relationships between key variables of interest (1993-2004)

<u>Share of women vs. Key firm characteristics</u>

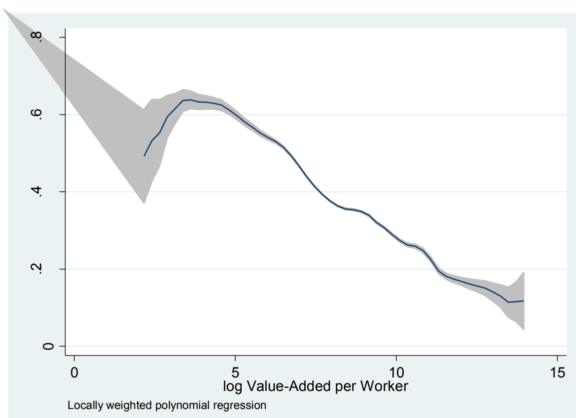
F1A: Gender composition vs. Firm-size



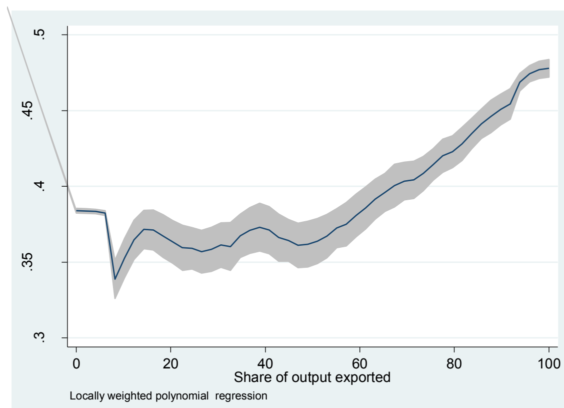
F1B: Gender Composition vs. Capital Intensity



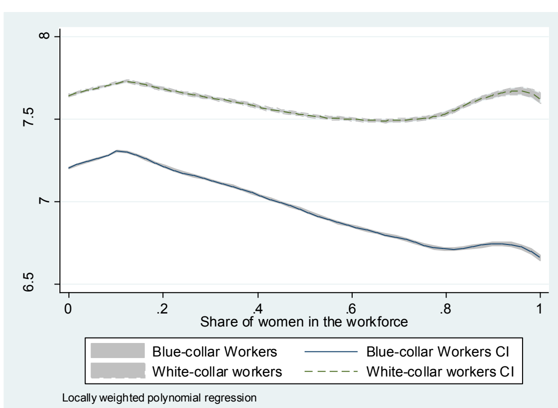
F1C: Gender composition vs. Value-added per Worker



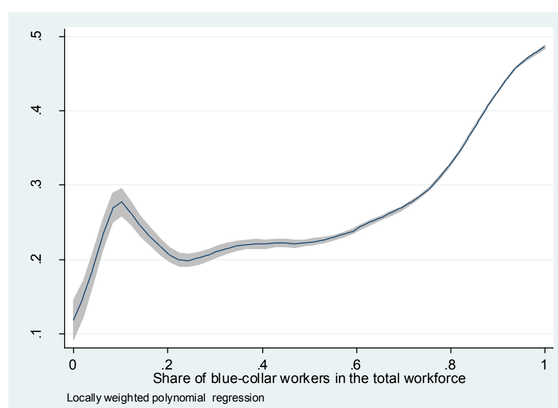
F1D: Gender Composition vs. Output Exported



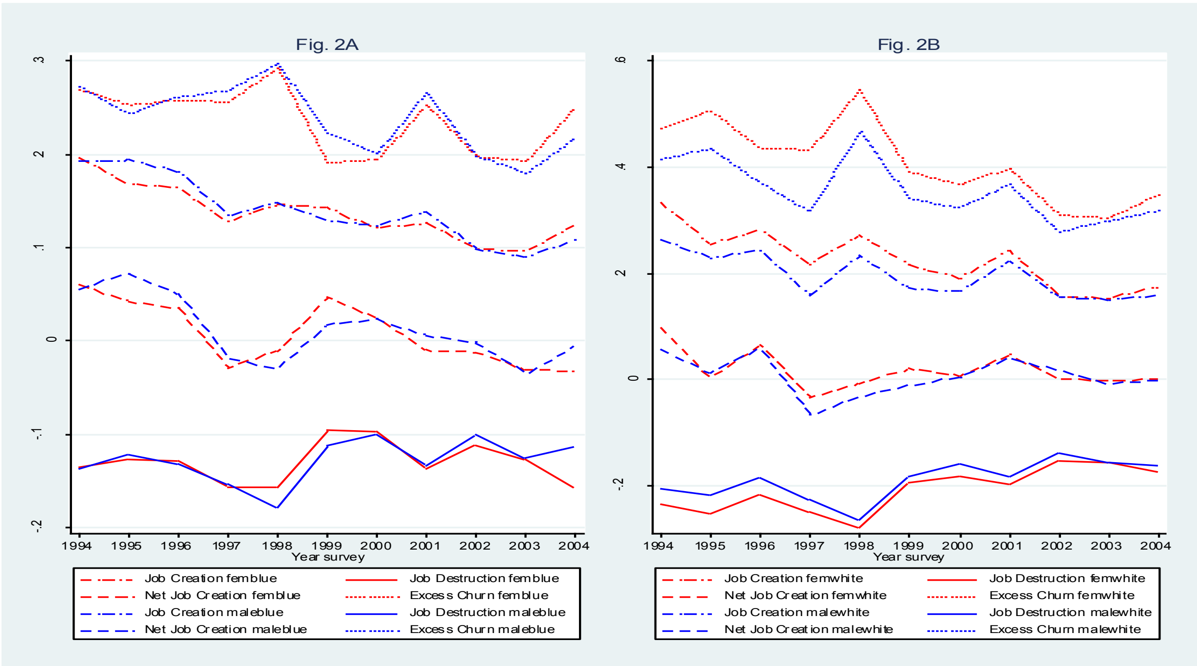
F1E: Gender composition vs. Wages



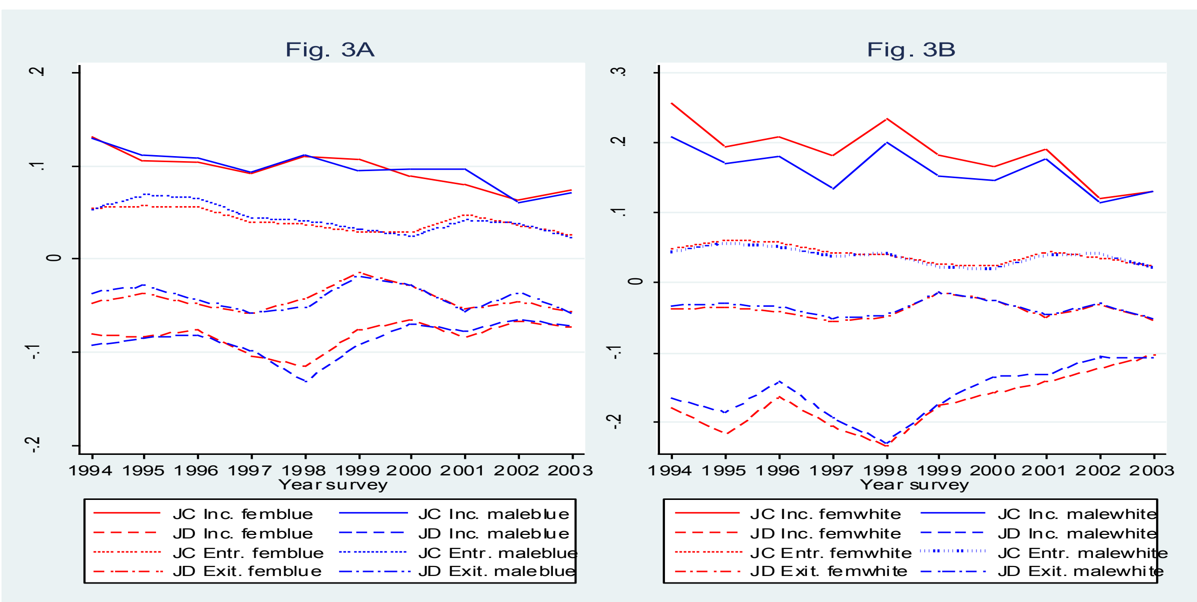
F1F: Gender composition vs. Unskilled Ratio



Figures 2A and 2B: Job Creation and Destruction, Net Job Creation, and Excess Churning (1994-2004)



Figures 3A and 3B: Job Creation and Destruction: Incumbents, Exiters and Entrants (1994-2003)



Wage Adjustment

Fig. 4A: Wage adjustment blue-collar workers – SI data

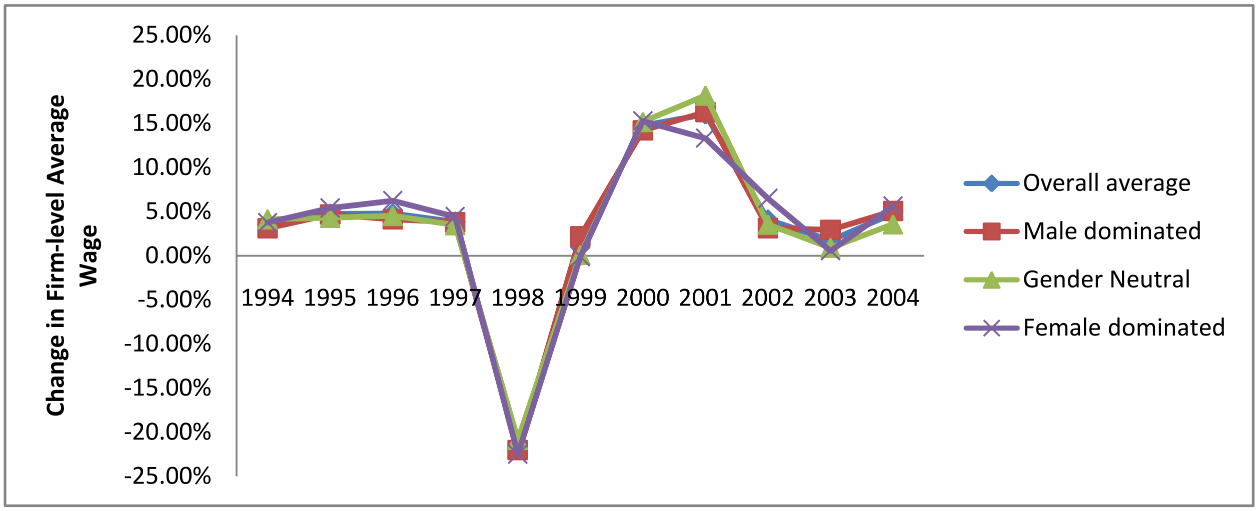


Fig. 4B: Wage adjustment white-collar workers – SI data

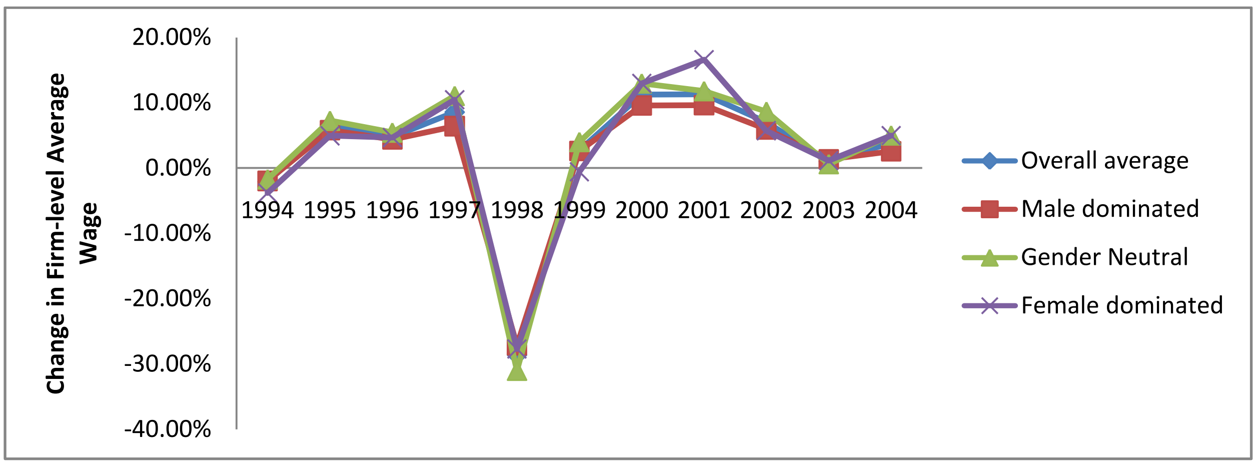


Fig. 4C: The Evolution of Inflation

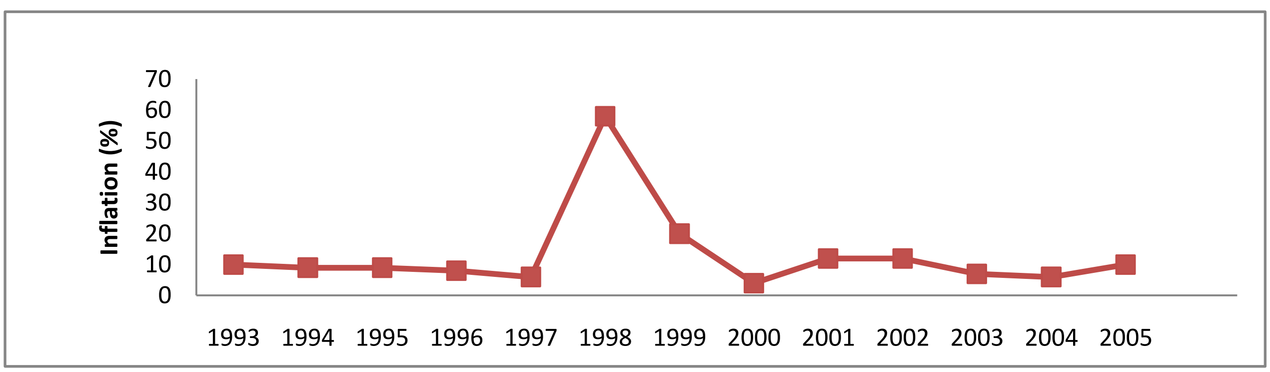


Fig. 5A: Entry and Exit Rates

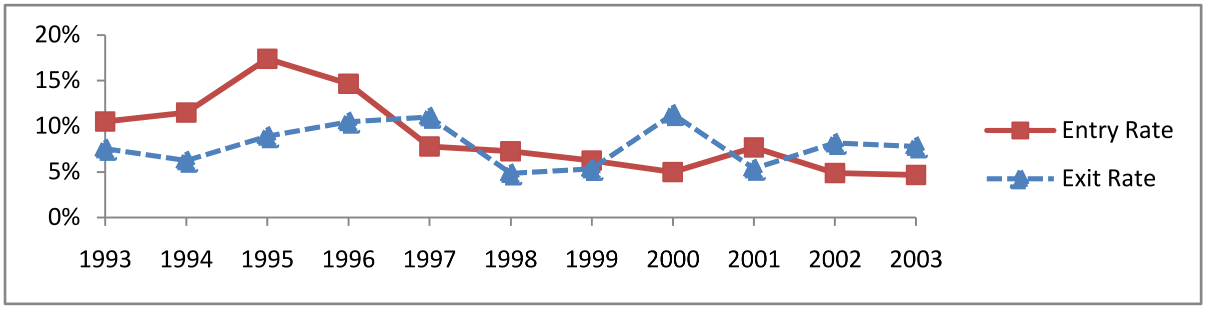
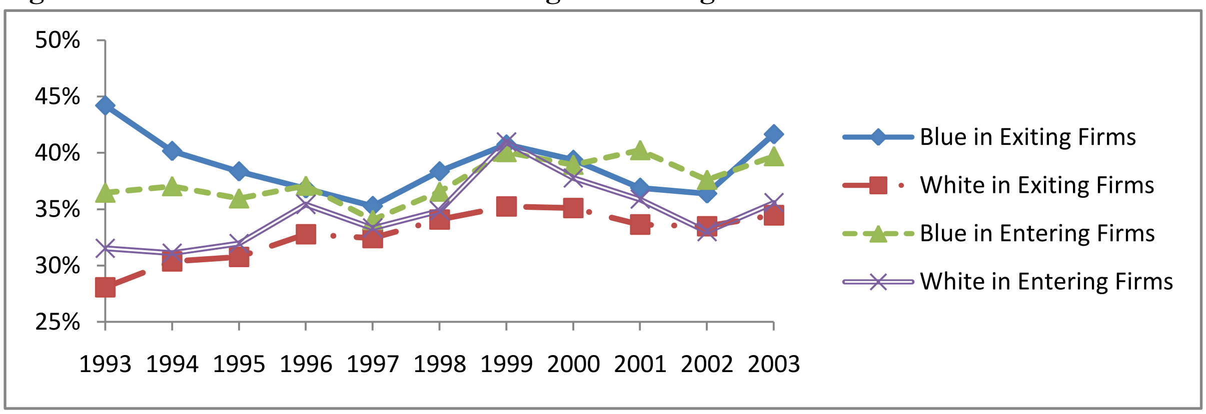


Fig. 5B: Share of female workers in Entering and Exiting Firms



Figures 6A and 6B: Employment growth vs the gender composition of the workforce (1993-2004)

Figure 6A: Blue-collar workers

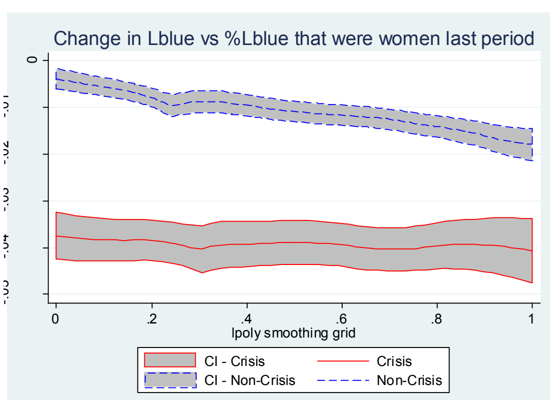
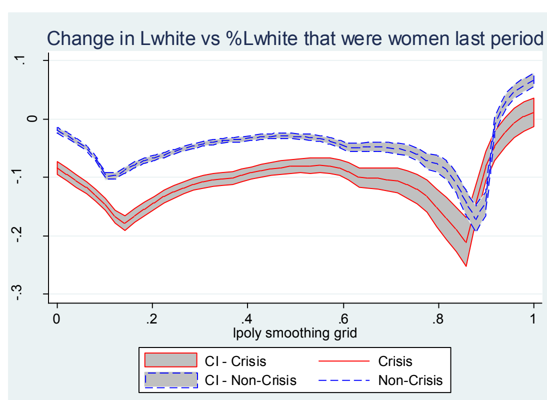
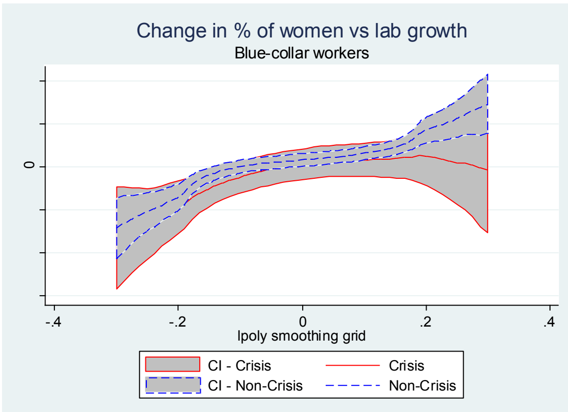


Figure 6B: White-collar workers



Figures 7A and 7B: Change in the gender composition of the workforce vs. emp. growth (1993-2004)



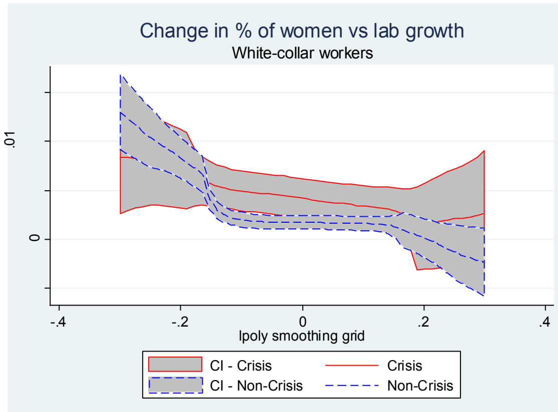


Table 1: Gender Composition by Industry (1993-2004)

	Share of workers that are Women	Share of blue-collar workers that are female	Share of white-collar workers that are female
Tobacco products	71.68%	75.43%	29.45%
Apparel	60.95%	62.48%	45.23%
Radio, television and communication equipment & medical instruments	50.04%	51.53%	38.30%
Textiles	49.50%	50.80%	36.62%
Leather	45.42%	45.37%	39.30%
Food products and beverages	41.26%	42.01%	28.76%
Recycling, rubber and plastics	38.07%	37.72%	35.45%
Coke, petroleum & chemical products	34.78%	34.88%	31.35%
Printing	31.51%	28.20%	41.34%
Furniture	31.31%	30.25%	35.56%
Paper products	30.96%	29.29%	35.01%
Electrical machinery & office equipment	29.41%	27.82%	31.51%
Other non-metallic mineral products	26.96%	26.22%	22.30%
Wood	23.08%	21.96%	29.58%
Basic metals & fabricated metal products, except machinery	19.24%	16.29%	31.13%
Other transport equipment	14.77%	11.44%	31.69%
Machinery and equipment	14.42%	10.63%	31.92%
Motor vehicles, trailers and semi-trailers	11.03%	7.27%	29.48%

Table 2: Descriptive Statistics Firms (N=157,495)

Variable	Explanation	Mean	sd	Weighted Mean
Share of female workers		0.37	0.29	0.46
%Women Blue Collar	Share of blue-collar workers who are female	0.37	0.32	0.47
%Women White Collar	Share of white-collar workers who are female	0.33	0.27	0.33
LogWblue	Log firm average blue-collar worker wage '000s Rupiah in 1993 prices	7.04	0.72	7.27
LogWwhite	Log firm-average white-collar worker wage '000s Rupiah in 1993 prices	7.53	0.95	8.00
LogL	Log total employment (blue-collar workers + white-collar workers)	4.06	1.19	5.99
LogLblue	Log blue-collar workers	2.20	1.41	4.01
LogLwhite	Log white-collar workers	4.27	1.16	6.20
Log(firm age+1)		2.30	0.86	2.43
Government owned	=1 if local or national government has owns any part of the company	0.03	0.16	0.06
Foreign owned	(=1 if any for. ownership)	0.07	0.25	0.19
Exporter	(=1 if any output exported)	0.18	0.38	0.40
Unskilled Ratio	(= blue-collar workers/tot. emp.)	0.82	0.15	0.84
Log(V/L)	= log (Real value-added per worker)	8.09	1.14	8.60
Log(K/L)	= log capital per worker measured as the deflated replacement value of machinery and equipment in '000 rupiah in 1993 prices	6.77	1.84	7.58
Log Minimum Wage	=log provincial level minimum wage in '000 rupiah in 1993 prices	6.92	0.25	6.93

Table 3: Firm Exit (1993-2003)

	Firm Exit: Dependent Variable: Log hazard Ratio (1)	(2)	(3)	(4)
	"Baseline"	**"Full Model"**	**"Baseline"** Disaggregated by occupation	**"Full Model"** Disaggregated by occupation
	coef/se	**coef/se**	**coef/se**	**coef/se**
%Women Overall	-0.073*	0.032		
	(0.040)	(0.048)		
Crisis*%Women Overall	-0.324***	-0.065		
	(0.091)	(0.106)		
%Women Blue			-0.064	0.073
			(0.040)	(0.047)
Crisis*%Women Blue			-0.255***	0.050
			(0.074)	(0.088)
%Women White			-0.090*	-0.100**
			(0.046)	(0.044)
Crisis*%Women White			0.134	0.009
			(0.082)	(0.079)
Unskilled ratio		-0.235**	0.442***	-0.252***
		(0.092)	(0.090)	(0.093)
Crisis*Unskilled ratio		0.052	-0.128	0.031
		(0.166)	(0.160)	(0.170)
LogL		-0.501***		-0.504***
		(0.016)		(0.016)
Crisis*LogL		-0.069**		-0.074**
		(0.030)		(0.031)
Log(V/L)		-0.169***		-0.167***
		(0.014)		(0.014)
Crisis*Log(V/L)		0.126***		0.132***
		(0.025)		(0.026)
Firmage (log)		-0.131***		-0.131***
		(0.015)		(0.015)
Crisis*Firmage(log)		-0.131***		-0.131***
		(0.026)		(0.026)
Foreign Owned		0.192***		0.189***
		(0.065)		(0.065)
Crisis*Foreign Owned		-0.265**		-0.269**
		(0.125)		(0.125)
Government Owned		0.214**		0.213**
		(0.094)		(0.094)
Crisis*Government Owned		-0.043		-0.035
		(0.179)		(0.180)
Exporter		0.199***		0.199***
		(0.039)		(0.039)
Crisis*exporter		-0.334***		-0.338***
		(0.076)		(0.076)
F-Test: Unskilled ratio+			*5.55*	*9.76*
*Crisis *Unskilled ratio=0*			*(p=0.02)*	*(p=0.00)*
F-Test: %Women Blue Collar			*25.69*	*5.11*
+crisis%Women Blue Collar=0			*(p=0.00)*	*(p=0.08)*
F-Test: %Women White Collar			*0.42*	*7.08*
+crisis%Women White Collar=0			*(p-0,52)*	*(p=0.03)*
Year dummies	***X***		***X***	
*Sector*Year Dummies*		***X***		***X***
Province Dummies		***X***		***X***
Number of observations	157,495	157,495	157,495	157,495
Pseudo R2	0.024	0.069	0.024	0.069

note: *** p<.01, ** p<.05, * p<.1

Table 4: Labor Growth (1993-2004)

	Labor growth							
	Blue-Collar Workers				**White Collar Workers**			
Dependent Variable	**Haltiwanger Growth Blue-collar Workers**				**Haltiwanger Growth White-collar Workers**			
Haltiwanger Growth $H=(L_t - L_{t-1})/0.5(L_t + L_{t-1})$	**(1) OLS coef/se**	**(2) FE coef/se**	**(3) OLS coef/se**	**(4) FE coef/se**	**(5) OLS coef/se**	**(6) FE coef/se**	**(7) OLS coef/se**	**(8) FE coef/se**
%Women blue $_{(t-1)}$	-0.000	-0.042***	0.021***	0.017**				
	(0.002)	(0.010)	(0.003)	(0.008)				
Crisis* %Women blue $_{(t-1)}$	0.019***	0.018***	0.012*	0.018***				
	(0.006)	(0.006)	(0.006)	(0.006)				
%Women white $_{(t-1)}$					0.040***	0.098***	0.035***	0.054***
					(0.005)	(0.012)	(0.005)	(0.010)
Crisis*%Women white $_{(t-1)}$					-0.004	-0.005	-0.021*	-0.021*
					(0.013)	(0.013)	(0.012)	(0.011)
LogLblue $_{(t-1)}$			-0.021***	-0.137***				
			(0.001)	(0.004)				
Crisis*LogLblue $_{(t-1)}$			-0.005**	-0.009***				
			(0.002)	(0.003)				
ΔLogWblue			-0.033***	-0.036***				
			(0.002)	(0.002)				
Crisis***Δ**LogWblue			-0.015***	-0.012**				
			(0.006)	(0.005)				
LogWblue $_{(t-1)}$			-0.012***	-0.027***				
			(0.001)	(0.002)				
Crisis*LogWblue $_{(t-1)}$			-0.002	-0.000				
			(0.003)	(0.003)				
LogLwhite $_{(t-1)}$							-0.025***	-0.097***
							(0.002)	(0.005)
Crisis*LogLwhite $_{(t-1)}$							-0.003	-0.007**
							(0.004)	(
ΔLogWwhite							-0.052***	-0.060***
							(0.002)	(0.003)
Crisis***Δ**LogWwhite							-0.022***	-0.022***
							(0.006)	(0.006)
LogWwhite $_{(t-1)}$							-0.010***	-0.038***
							(0.001)	(0.003)
Crisis*LogWwhite $_{(t-1)}$							-0.004	-0.004
							(0.003)	(0.003)
ΔLogY			0.047***	0.046***			0.034***	0.033***
			(0.002)	(0.002)			(0.002)	(0.002)
Crisis***Δ**LogY			0.013***	0.014***			0.010*	0.014**
			(0.004)	(0.004)			(0.005)	(0.005)
LogY $_{(t-1)}$			0.015***	0.035***			0.017***	0.028***
			(0.001)	(0.002)			(0.001)	(0.002)
Crisis*LogY $_{(t-1)}$			0.000	0.002			0.000	0.003
			(0.002)	(0.002)			(0.003)	(0.002)
Foreign owned $_{(t-1)}$			0.001	-0.002			-0.009***	-0.000
			(0.002)	(0.005)			(0.003)	(0.007)
Crisis* Foreign owned $_{(t-1)}$			-0.000	-0.001			0.023***	0.021***
			(0.005)	(0.005)			(0.007)	(0.007)
Government owned $_{(t-1)}$			-0.007**	-0.011**			0.009*	-0.011
			(0.003)	(0.006)			(0.005)	(0.008)
Crisis* Government owned			0.003	-0.001			0.022*	0.036***
			(0.008)	(0.009)			(0.011)	(0.010)
Exporter $_{(t-1)}$			0.002	0.004			0.001	0.002
			(0.002)	(0.002)			(0.003)	(0.003)
Crisis*exporter $_{(t-1)}$			0.009**	0.004			0.001	-0.005
			(0.004)	(0.004)			(0.006)	(0.006)
Unskilled ratio $_{(t-1)}$			-0.064***	-0.064***			0.045***	0.125***
			(0.007)	(0.015)			(0.008)	(0.023)
Crisis*Unskilled ratio $_{(t-1)}$			-0.024	-0.043***			0.034*	0.031*
			(0.015)	(0.014)			(0.018)	(0.019)
Firmage (log) $_{(t-1)}$			-0.006***	-0.003			-0.004**	-0.008
			(0.001)	(0.004)			(0.001)	(0.005)
Crisis*Firmage (log) $_{(t-1)}$			-0.004*	-0.001			0.004	0.002
			(0.002)	(0.002)			(0.003)	(0.003)
Year Dummies	***X***	***X***			***X***	***X***		
Sector-Year Dummies			***X***	***X***			***X***	***X***
Province Dummies			***X***				***X***	
*F-test: Gender-composition +crisis*gender composition=0*	*13.76 (p=0.00)*	*5.04 (p=0.03)*	*34.51 (p=0.00)*	*12.91 (p=0.00)*	*9.30 (p=0.00)*	*35.70 (p=0.00)*	*1.74 (p=0.19)*	*6.18 (p=0.01)*
Number of observations	124.536	124.536	124.536	124.536	124.536	124.536	124.536	124.536
R2	0.009	0.022	0.124	0.248	0.006	0.013	0.130	0.217
Adjusted R2	0.009	0.022	0.122	0.247	0.006	0.013	0.129	0.216

Notes:

*** $p<0.01$, ** $p<0.05$, * $p<0.1$

The Haltiwanger growth measure is defined as $H=(L_t - L_{t-1})/0.5(L_t + L_{t-1})$

Standard errors are heteroscedasticity robust and clustered by establishment.

The sample is restricted to firms for which the absolute change in employment was lower than 100%.

OLS Weighted regressions use as weight the average number of blue-collar workers over the period over which the growth is defined, FE Weighted regressions use as weights the average number of blue-collar workers over the entire period

Table 5: Differential treatment

	Differential Treatment							
	Blue-Collar Workers				**White Collar Workers**			
Dependent Variable	**Change in the share of blue-collar workers that are women**				**Change in the share of white-collar workers that are women**			
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**	**(8)**
	OLS	**FE**	**OLS**	**FE**	**OLS**	**FE**	**OLS**	**FE**
	coef/se	**coef/se**	**coef/se**	**coef/se**	**coef/se**	**coef/se**	**coef/se**	**coef/se**
***Δ*LogLblue- positive**	0.008**	0.008*	0.008**	0.007***				
	(0.004)	(0.004)	(0.004)	(0.002)				
***Δ*LogLblue - negative**	0.009*	0.009	0.009*	0.009***				
	(0.005)	(0.005)	(0.005)	(0.002)				
Crisis****Δ*LogLblue-positive**	0.007	0.006	0.008	0.007*				
	(0.009)	(0.009)	(0.009)	(0.003)				
Crisis****Δ*LogLblue -negative**	-0.000	0.007	-0.001	0.007**				
	(0.009)	(0.009)	(0.009)	(0.003)				
***Δ*LogLwhite- positive**					-0.040***	-0.043***	-0.041***	-0.044***
					(0.005)	(0.005)	(0.005)	(0.002)
***Δ*LogLwhite- negative**					-0.038***	-0.042***	-0.039***	-0.042***
					(0.005)	(0.006)	(0.005)	(0.002)
Crisis****Δ*LogLwhite- positive**					0.018*	0.018*	0.017*	0.018***
					(0.010)	(0.011)	(0.010)	(0.004)
Crisis****Δ*LogLwhite-negative**					-0.006	-0.010	-0.004	-0.008**
					(0.010)	(0.011)	(0.010)	(0.003)
Lblue (log) $_{(t-1)}$			-0.001***	-0.003***				
			(0.000)	(0.001)				
crisis*Lblue (log) $_{(t-1)}$			-0.000	-0.001				
			(0.001)	(0.000)				
Lwhite (log) $_{(t-1)}$							-0.001*	-0.003***
							(0.000)	(0.001)
crisis*Lwhite (log) $_{(t-1)}$							-0.000	-0.000
							(0.001)	(0.001)
Foreign Owned $_{(t-1)}$			0.001	0.001			0.001	0.001
			(0.001)	(0.001)			(0.002)	(0.002)
Crisis*Foreign Owned $_{(t-1)}$			-0.001	0.001			0.001	0.000
			(0.003)	(0.001)			(0.004)	(0.002)
Government Owned $_{(t-1)}$			0.001	-0.004**			-0.001	-0.001
			(0.001)	(0.002)			(0.002)	(0.002)
Crisis*Government $_{(t-1)}$ Owned			0.005	0.004			-0.005	-0.006***
			(0.003)	(0.002)			(0.005)	(0.002)
Exporter $_{(t-1)}$			0.001	0.001**			0.002*	0.002**
			(0.001)	(0.001)			(0.001)	(0.001)
Crisis*exporter $_{(t-1)}$			0.001	0.001			-0.004	-0.005***
			(0.002)	(0.001)			(0.003)	(0.001)
Unskilled ratio $_{(t-1)}$			0.008***	0.024***			-0.001	-0.017***
			(0.002)	(0.003)			(0.003)	(0.004)
Crisis*Unskilled ratio $_{(t-1)}$			-0.002	-0.005			-0.008	-0.008*
			(0.007)	(0.004)			(0.008)	(0.004)
Firmage (log) $_{(t-1)}$			-0.000	-0.003**			0.000	-0.002
			(0.000)	(0.001)			(0.001)	(0.002)
Crisis*Firmage(log) $_{(t-1)}$			0.001	0.002***			0.000	0.001
			(0.001)	(0.001)			(0.002)	(0.001)
*F-test: **Δ**LogL-positive =*	*0.03*	*0.02*	*0,04*	*0,58*	*0.08*	*0.05*	*0.14*	*1.12*
***Δ**LogL negative*	*(p=0.87)*	*(p=0.88)*	*(p=0.84)*	*(p=0.45)*	*(p=0.77)*	*(p=0.82)*	*(p=0.71)*	*(p=0.29)*
*F-test: **Δ**LogL-negative*	*3.42*	*2.53*	*3.48*	*18.75*	*5.82*	*0.65*	*7.07*	*72.32*
*+Crisis***Δ**LogL-negative=0*	*(p=0.06)*	*(p=0.11)*	*(p=00.6))*	*(p=0.00)*	*(p=0.02)*	*(p=0.01)*	*(p=0.00)*	*(p=0.00)*
*F-test: **Δ**LogL-positive*	*1.31*	*4.24*	*1.09*	*30.39*	*26.74*	*32.25*	*7.07*	*277.67*
*+Crisis***Δ**LogL-positive=0*	*(p=0.25)*	*(p=0.04)*	*(p=0.30)*	*(p=0.30)*	*(p=0.00)*	*(p=0.01)*	*(p=0.00)*	*(p=0.00)*
Year Dummies	***X***	***X***			***X***	***X***		
Sector-Year Dummies			***X***	***X***			***X***	***X***
Province Dummies			***X***				***X***	
Number of observations	119,818	119,818	119,818	119,818	119,818	119,818	119,818	119,818
R2	0.002	0.002	0.011	0.012	0.021	0.024	0.031	0.034
Adjusted R2	0.002	0.002	0.009	-0.250	0.021	0.024	0.029	-0.222

Notes:
*** p<0.01, ** p<0.05, * p<0.1
Standard errors are heteroscedasticity robust and clustered by establishment
OLS Weighted regressions use as weight the average number of blue-collar workers over the period over which the growth is defined, FE Weighted regressions use as weights the average number of blue-collar workers over the entire period

Table 6: The Evolution of Gender Wage Gaps

	Firm-level Earnings Regressions							
	Blue-Collar Workers				White Collar Workers			
	Dependent Variable				Dependent Variable			
	Log average wages blue-collar workers				Log average wages white-collar workers			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	OLS	FE	OLS	FE	OLS	FE	OLS	FE
	coef/se	coef/se	coef/se	coef/se	coef/se	coef/se	coef/se	coef/se
%Women Blue	-0.358***	-0.176***	-0.144***	-0.069				
	(0.029)	(0.051)	(0.024)	(0.045)				
Crisis* %Women Blue	0.007	-0.021	-0.023	-0.033				
	(0.029)	(0.022)	(0.032)	(0.026)				
%Women White					-0.215***	0.141**	-0.163***	0.008
					(0.082)	(0.062)	(0.062)	(0.049)
Crisis* %Women White					0.070	0.053	0.045	0.007
					(0.074)	(0.054)	(0.062)	(0.049)
Unskilled Ratio			-0.799***	-1.007***			1.114***	1.932***
			(0.051)	(0.064)			(0.060)	(0.099)
Crisis*Unskilled Ratio			0.127*	0.224***			0.257***	0.206***
			(0.074)	(0.064)			(0.086)	(0.068)
Log L			0.053***	-0.070***			0.043***	-0.138***
			(0.006)	(0.017)			(0.010)	(0.028)
Crisis*Log L			-0.014*	-0.017***			-0.001	0.000
			(0.008)	(0.007)			(0.012)	(0.009)
Log (V/L)			0.204***	0.122***			0.259***	0.111***
			(0.007)	(0.007)			(0.010)	(0.010)
Crisis*Log (V/L)			0.017*	0.019**			-0.004	-0.007
			(0.009)	(0.008)			(0.014)	(0.011)
Foreign Owned			0.070***	0.011			0.206***	0.020
			(0.021)	(0.030)			(0.034)	(0.056)
Crisis*Foreign Owned			0.022	0.001			0.021	0.044
			(0.028)	(0.025)			(0.038)	(0.032)
Government Owned			0.035	0.009			-0.043	0.041
			(0.036)	(0.041)			(0.038)	(0.046)
Crisis*Government			-0.005	-0.001			0.033	0.025
			(0.046)	(0.043)			(0.060)	(0.056)
Exporter			0.063***	0.028**			0.126***	0.020
			(0.013)	(0.013)			(0.021)	(0.018)
Crisis*exporter			-0.023	-0.028			-0.063	-0.043
			(0.026)	(0.021)			(0.040)	(0.027)
Firm age (log)			0.018*	0.103***			0.010	0.117***
			(0.010)	(0.026)			(0.017)	(0.043)
Crisis*Firmage (log)			0.017	0.010			0.034*	0.008
			(0.012)	(0.010)			(0.019)	(0.015)
Log Minimum Wage			0.022	0.037			-0.021	-0.015
			(0.050)	(0.045)			(0.079)	(0.073)
Crisis*Log Minimum			-0.164**	-0.111*			0.019	0.038
			(0.065)	(0.058)			(0.105)	(0.088)
Year Dummies	***X***	***X***			***X***	***X***		
Sector-Year Dummies			***X***	***X***			***X***	***X***
Province Dummies			***X***				***X***	
N	138,133	138,133	138,133	138,133	138,133	138,133	138,133	138,133
R2	0.059	0.107	0.366	0.184	0.010	0.028	0.311	0.156
Adjusted R2	0.059	0.107	0.365	0.183	0.010	0.028	0.310	0.154

Notes:

*** p<0.01, ** p<0.05, * p<0.1

Standard errors are heteroscedasticity robust and clustered by establishment

OLS Weighted regressions use as weight the average number of blue-collar workers over the period over which the growth is defined, FE Weighted regressions use as weights the average number of blue-collar workers over the entire period

Appendix A: Examining the representativeness of our findings using SAKERNAS

This appendix contains an analysis of gender differences in the impact of the East Asian crisis on manufacturing wages and employment using SAKERNAS – a nationally representative labor force survey. This analysis complements the firm-level analysis presented in the main body of the text and helps assess the representativeness of the results obtained.

The SAKERNAS data contain detailed information on individual's gender, age, education, occupation and earnings. Yet, we cannot match individual workers to firms in the census. Moreover, while SAKERNAS enables us to identify which individuals are employed in manufacturing, information on firm-size is lacking. Thus the sample of manufacturing workers in SAKERNAS contains not only employees of firms observed in the SI data (that is, firms with more than 20 employees), but also many individuals who work in very small manufacturing firms, allowing a comparison of the entire manufacturing sector with the larger, more formal firms covered by the manufacturing census.

A.1 Employment adjustment

Figure A1 depicts the evolution of aggregate employment and manufacturing employment, separately for men and women. The crises only induced a small change in overall employment (consistent with the patterns discussed in Sections 2 and 5) and the differences between men and women were not large. Post-crisis, aggregate employment rates dropped, but by slightly more for women than for men. The evolution of manufacturing employment is more volatile and it appears as though the crisis induced a substantial reduction in manufacturing employment, and that this reduction was slightly larger for men than for women. Post-crisis, manufacturing employment rebounded, and men's manufacturing employment grew more quickly than women's. Note that these trends are not strictly comparable to those observed using the SI-data, which do not demonstrate an increase in the number of manufacturing jobs. Part of the reason might have to do with the fact that the SI data only covers large manufacturing firms; it may be the case that the number of small manufacturing firms has risen rapidly (perhaps partly reflecting an increase in informality in response to decreasing labor market flexibility due to more stringent labor laws).

A.2 Wage Adjustment

Figure A2 depicts the evolution of average wages and average manufacturing wages for men and women separately using SAKERNAS data. These trends are consistent with those observed using the SI-data as both men's and women's real wages fell precipitously during the crisis in manufacturing and in other sectors. However, in contrast with the findings obtained using the SI data, men's average real manufacturing earnings appear to have fallen more than women's.

To assess whether the gender wage gap indeed closed, we estimate standard Mincerian equations that mimic the firm-level specifications presented in Section 7. That is we estimate equations of the form:

$$\mathrm{Log}(\mathrm{w}_i) = \varphi_{\mathrm{f}} fem_i + \varphi_{\mathrm{Crisis*f}}\mathrm{Crisis} * fem_i + \varphi_{\mathrm{P}} P_i + \varphi_{\mathrm{Crisis*P}}\mathrm{Crisis} * P_i + o_i$$

Where fem_i is a dummy variable indicating the gender of the individual and P_i is a vector of individual-specific characteristics, include their age and education, and o_i is an error term. φ_{f} provides a measure of the gender wage gap, while $\varphi_{\mathrm{Crisis*f}}$ tells us by how much the gender wage gap changed during the crisis. These equations are estimated separately for blue-collar and white-collar workers.

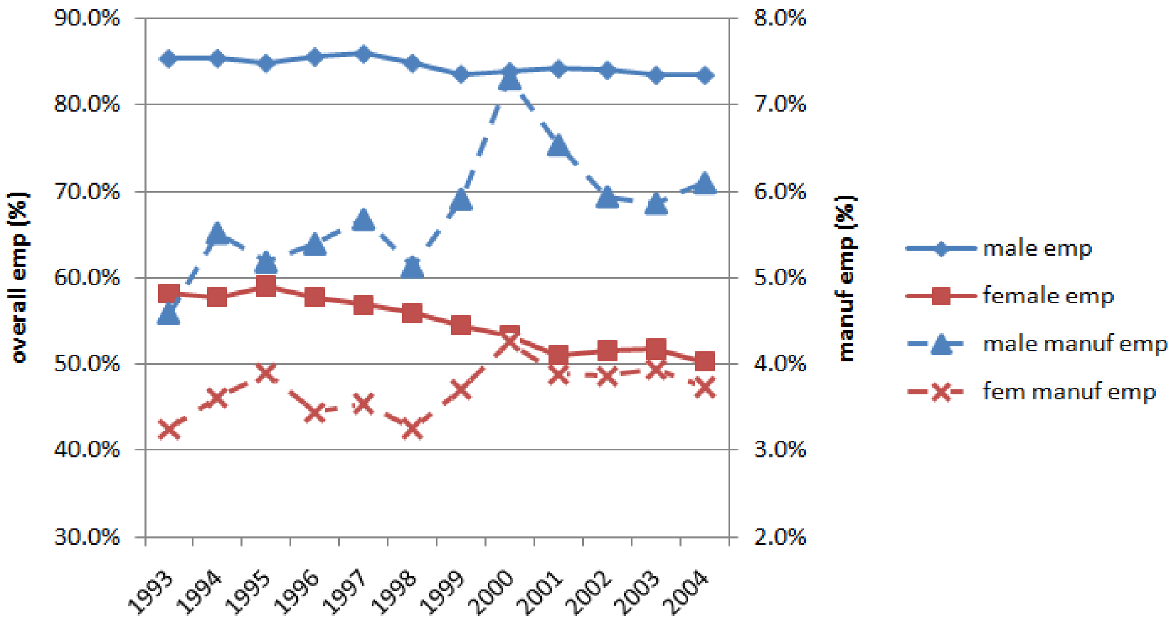
The results are presented in Table A1. The dependent variable is real monthly earnings. Column 1 only includes a gender dummy and its interaction with the crisis dummy and a control for the number of hours worked, as well as province year dummies. Column 2 adds age, age squared and dummies for educational attainment. Column 3 interacts these with crisis dummies. Columns 4, 5 and 6 repeat specifications 1, 2 and 3 respectively, for white-collar workers.

The raw gender premium in manufacturing is roughly 28% for blue-collar workers and 30% for white-collar workers. Controlling for age and education reduces these premia to 19% and 21% respectively. Note that these premia are similar to the ones obtained using the firm-level average wages (see Section 7). The crisis-gender interaction dummy is negative but never statistically significant for blue-collar workers, suggesting that women suffered more adverse wage losses even though we cannot reject the null that the gender wage gap did not shrink during the crisis (note that that the estimates imply that women's wages dropped by 4.9% more than men's wages; this number is strikingly similar to the estimates of the widening of gender pay gaps by approximately 4% obtained with the firm-level data). For white-collar workers it is positive and significant if education and age are not controlled for. Once individual characteristics are controlled for, however, the interaction of gender with the crisis dummy becomes insignificant. Thus, the null hypothesis that the gender-wage gaps for both blue-collar and white-collar workers did not change during the crisis cannot be rejected.

These results are consistent with those obtained using the SI data. While the gender wage gap in the SI increased on average, the increase in the wage gap was positively correlated with firm size. Women in the smallest SI firms experienced a reduction in the gender wage gap. Consequently, it should perhaps not come as a surprise that once we include employees of firms with fewer than 20 employees in our sample, the average wage gap is reduced.

Overall, then, the results obtained using the SAKERNAS dataset corroborate the results obtained using the SI data.

Graph A1: The Evolution of Aggregate Employment and Manufacturing Employment



Graph A2: The Evolution of Monthly Manufacturing Earnings

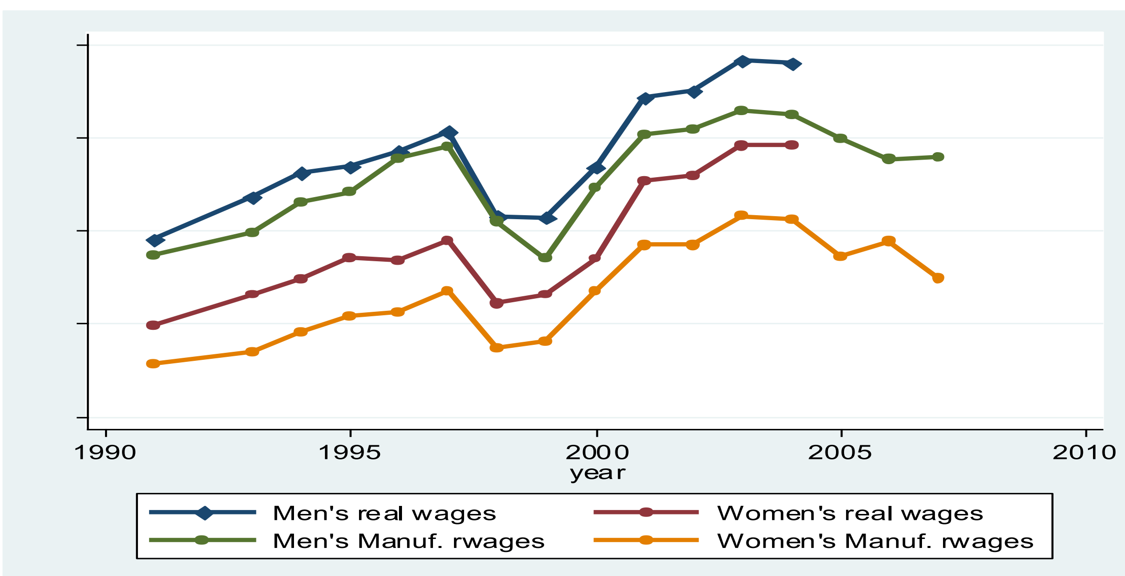


Table A1 – Earnings Regressions based on SAKERNAS Data

	Minercian Earnings Regressions (OLS) Dependent variable: log real monthly income					
	Blue-collar Workers			White-collar Workers		
	1	**2**	**3**	**4**	**5**	**6**
	coef/se	**coef/se**	**coef/se**	**coef/se**	**coef/se**	**coef/se**
Female	-0.284***	-0.187***	-0.188***	-0.304***	-0.217***	-0.213***
	(0.032)	(0.030)	(0.031)	(0.029)	(0.025)	(0.025)
Crisis*female	-0.049	-0.047	-0.046	0.066*	0.055	0.037
	(0.037)	(0.030)	(0.031)	(0.039)	(0.033)	(0.030)
Hours(log)	0.574***	0.542***	0.546***	0.362***	0.345***	0.326***
	(0.047)	(0.042)	(0.049)	(0.073)	(0.056)	(0.046)
Age		0.039***	0.039***		0.047***	0.050***
		(0.003)	(0.004)		(0.004)	(0.004)
Age^2		-0.000***	-0.000***		-0.000***	-0.000***
		(0.000)	(0.000)		(0.000)	(0.000)
Primary		0.257***	0.261***		0.294***	0.309***
		(0.018)	(0.026)		(0.023)	(0.027)
Secondary		0.474***	0.471***		0.593***	0.591***
		(0.025)	(0.031)		(0.027)	(0.029)
Tertiary		0.885***	0.919***		1.146***	1.134***
		(0.078)	(0.069)		(0.035)	(0.036)
Crisis*Hours(log)			-0.014			0.080
			(0.059)			(0.126)
Crisis*Age			0.001			-0.011
			(0.008)			(0.011)
Crisis*Age^2			-0.000			0.000
			(0.000)			(0.000)
Crisis*Primary			-0.015			-0.066
			(0.061)			(0.075)
Crisis*Secondary			0.012			0.012
			(0.064)			(0.058)
Crisis*Tertiary			-0.154			0.081
			(0.178)			(0.096)
Province Dummies	***X***	***X***	***X***	***X***	***X***	***X***
Year Dummies	***X***	***X***	***X***	***X***	***X***	***X***
Number of observations	21,883	21,883	21,883	20,162	20,162	20,162
R2	0.354	0.442	0.443	0.268	0.470	0.471
Adjusted R2	0.353	0.441	0.441	0.266	0.469	0.470

Notes:
*** p<0.01, ** p<0.05, * p<0.1, Robust standard errors are clustered by kabupaten/walikota (district/town). Regressions are weighted by sample weights unless otherwise noted. Sample is restricted to actively working manufacturing wage earners between ages of 15 and 70. Reference category for education dummies is no or uncompleted primary education.

Appendix B: Data Appendix

B.1 Construction of panel

Survei Industri (SI) from Indonesia's statistical bureau (*Badan Pusat Statistik—BPS)* provides the plant-level data used in this paper. The SI is an annual census of all manufacturing establishments with more than 20 employees.[36] While the annual survey has been conducted going back to 1975, field procedures for identifying new firms were dramatically improved over 1985-1990, making the accuracy of measuring entry prior to 1990 problematic (see Aswicahyono, 2008). Moreover, the dataset does not consistently differentiate male versus female blue-collar or white-collar workers prior to 1993, so as a result, this paper only considers the data post 1993. We use data up until 2004, since data from later years are not immediately comparable due to changes in firm identifiers, changing definitions of key explanatory variables, and province splits.

B.2 Construction of Key Explanatory Variables

Sector

Sector of main product In order to classify establishments by industry, BPS records the five-digit International Standard Industrial Classification (ISIC) for firms based on the product with the largest production value in any given year. In 2001, BPS changed the classification of plants from the second revision of the ISIC to the third. A consistent bridge between these different coding systems was constructed based upon the inclusion of both codings in the dataset for the 2000 survey database. This bridge was corroborated using a bridge provided by BPS. In many cases, the industry code provided in the dataset was truncated to four or fewer digits. Where possible, if an adjacent year's reported industry for the same firm was available that was used to fill in the truncated digits. Where plants produce multiple products or the production processes allow changing from one product to another, we may expect to see the coded industry of production change from year to year. In such cases, the mode sector code is used (with ties going to the initial sector code reported).

Labor

Total labor for an establishment was defined as the sum of all paid and non-paid workers, whether production or non-production workers.[37] Production workers were defined by BPS as all workers who work directly in the production process or activities connected with production process, and non-production workers were defined to be all other workers. These definitions roughly correspond to traditional definitions of blue- and white-collar workers, respectively (see below).

Blue collar workers the sum of all production and unpaid workers. After 2000, we are unable to distinguish paid from unpaid workers in production and non-production occupations, and prior to 2001, we are unable to distinguish whether unpaid workers are in production or non-production. As a result, we do not include them in the labor variables prior to 2001, but there is a slight jump in employment as a result of this discrepancy in this year.

[36] By law, all firms with over 20 employees are required to complete the BPS's questionnaire every year. Every ten years (1986, 1996, 2006) BPS devotes additional resources to track down all firms to ensure that any firms that had not been reporting do so. A major tracking effort was thus undertaken immediately prior to the crisis, and we can thus be reasonably confident that the data are not missing out on a substantial group of non-reporting firms.
[37] Although BPS does not explicitly define it, this is understood to include both permanent and temporary workers.

White collar workers: the sum of all non-production workers.

Unskilled ratio. The unskilled ratio was constructed as the share of blue collar workers (including unpaid production workers) as a share of the total establishment employment.

Share of female workers. The share of female workers (whether blue or white-collar workers) is constructed, quite simply, by dividing the number of female blue-(white-)collar workers by the total number of blue-(white-)collar workers.

The establishment-level **total wage bill** was constructed as the sum of cash wages/salary and in-kind benefits for blue and white-collar workers deflated to 1993 rupiah using the national consumer price index obtained from the World Development Indicators.

Wages. Wages are defined as the average wage for blue- or white collar workers, constructed as the total wage bill for either group divided by the number of workers of either respective group. Real wages are constructed by deflating the nominal wage bill reported to BPS by provincial CPI obtained from BPS.

Minimum wages. Minimum wages were obtained from BPS as monthly provincial minimum wages set by each province (or averages where there is within-province variation across districts) for each year. These reflect the nominal rupiah amount that all formal sector workers are required to be paid at or above. To obtain minimum wages in real terms, we deflated provincial minimum wages by the corresponding provincial CPI obtained from BPS.

Capital and materials

Capital was defined using the provided estimated value of machinery and equipment for each establishment in a given year. Where the estimated value was not available, the book value was used as provided. These values were deflated to 1993 rupiah values on an annual basis using the reported GDP deflator for machinery for 1993-2001 provided by BPS and using the manufacturing sector deflator for 1990-1992. The manufacturing survey does not include data on the value of the capital stock for 1996, so these values were interpolated using predicted values constructed from a regression of the capital stock on contemporaneous output, material inputs, labour usage, ownership characteristics, whether or not the firm exports, province and lagged capital for 1991-1995. In a separate imputation procedure, we also imputed capital for firms that did not report their capital stocks.

Materials were defined using the total reported value of raw materials used by the firm. These were deflated to 1993 rupiah using a deflator constructed from 2-digit industry GDP deflators weighted by the shares of input indicated from input-output tables obtained from the OECD Input-Output Table Database for 1995 and 2000.

Value-added and gross output

Value-added and gross output were calculated by BPS on the basis of the total value of production (and net of total expenditure for value added). Both variables were deflated to 1993 rupiah using a 2-digit industry level GDP deflator.

Firm age

Firm age was constructed using the difference between the survey year and the year the firm reported the start of production.

Export Status and ownership

Exporters were identified for 1993-2000 based upon reporting that the share of output exported was nonzero and non-missing, and for 2001-2004 based upon a variable identifying whether *any* output was exported (both variables are not available in the data for all years.)

Foreign owned firms based upon whether foreign ownership was reported to be non-missing and non-zero.

Government owned firms were defined the same way based on both local and national government.

Entry and exit

Entry is defined as the first year an establishment is observed in the data. Firms may have been in operation for several years prior to crossing the minimum threshold of 20 employees to be included in the survey. **Firm exit** is defined as permanent exit from the dataset. Note that, since the survey only includes firms with 20 or more employees, firms whose employment falls below this threshold will be defined as having exited in case they do not reappear in the data. There is a spike in exit rates in 2000 that is due partially to economic factors (a recession) and data issues that are likely to have been caused by the "big bang" decentralization of central government in Indonesia at that time and the splitting off of new provinces that occurred as a result. Because the statistics ministry manages its enumerators out of province offices, some anecdotal evidence points to the fact that tracking firms became more difficult and may have resulted in spurious exit (see Hallward-Driemeier and Rijkers, 2010 for a more detailed description).

B. 3 Cleaning of outliers

The problem of non-persistent extreme values is recognized in published work using the SI (see for example, Blalock, Gertler and Levine, 2008). In some cases, these values may be the result of key punch errors, where, for example, ownership share is recorded as 340 percent rather than 34 percent. Where shares (exports and ownership) were reported, these were easily corrected, but for balance sheet variables, more extensive cleaning was required.

These likely key punch errors and other highly volatile trends were corrected in the data by identifying firms with implausibly large non-persistent changes over one or two years and replacing values with those of preceding and succeeding years (or just one or the other of the observation was at beginning or end of the series). Based on close observation of normal variation of these variables, the thresholds for identifying large non-persistent shifts where other variables did not shift as dramatically were a 100% change in labor, a 200% in real value added and real output, a 150% change in real inputs, and 100% change in capital and average wages. Next, firms were dropped from the data when they had such a significant number of non-persistent jumps that it was impossible to interpolate (this constituted less than 1% of the sample). Finally remaining outliers were identified by eyeballing plots of key relationships (e.g. inputs per worker) and spot interpolating obvious remaining outliers.